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Mergers and Beliefs

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Abstract

We study the combined effects of managerial optimism and market overvaluation on merger premiums and the chosen form of payment. Our empirical results are consistent with market overvaluation and the target manager's optimism as having the most influence on mergers. The observed form of payment corresponds to the acquiring manager's preferences, suggesting that the acquiring manager dictates the method of payment. Lastly, our model demonstrates why cash mergers are more likely to be hostile, and provides an explanation for why a combination of cash plus stock may be optimal.

Introduction

Recent research has examined separately the effects of both managerial and market optimism on mergers. According to this view when the market is overvalued, companies will try to acquire other companies by using their overvalued stock as payment (e.g. Shleifer and Vishny, 2003 and Dong et al., 2006). Conversely, when CEO's are optimistic in relation to their ability to manage, they prefer to use cash to acquire other companies instead of their stock, which they perceive to be undervalued (e.g. Roll, 1986 and Malmendier and Tate, 2008). In this study we allow for both motives for mergers.

Many explanations for mergers and acquisitions have been based on the joint assumption of unbiased managerial behavior and efficient capital markets. In this context, mergers can be explained as a rational attempt to capture synergies or eliminate agency problems. In loosening the assumption about managerial behavior, Roll (1986) suggests that managerial hubris may also help explain observed merger activity. Managerial hubris causes managers overestimate the probability that they will operate the target company better than the current management, therefore causing them to overpay for the target. Shefrin (2001) also notes that psychological biases may prevent decision makers from operating in an unbiased manner.

Malmendier and Tate (2008) test the hubris hypothesis by sorting managers according to their perceived optimism. They define optimistic managers as those who consistently overestimate their ability in acquiring and running potential projects. They conclude that optimistic managers pay too much for target firms while misjudging their potential value-added, thus acquiring companies that could be value-destroying. Furthermore, optimistic CEO's view issuing equity as costly because they frequently believe that their firm is undervalued. This forces optimistic CEO's to use cash in purchasing other firms to avoid issuing new undervalued equity. This need for cash exacerbates the effects of overconfident managers with an abundance of internal resources and also denies the company some otherwise profitable opportunities (where equity could be used). Their empirical tests generally support these predictions.

Shleifer and Vishny (2003) relax the assumption of market efficiency, and suggest that mergers are a rational response by managers trying to time the market or take advantage of private information about the misvaluation of their own firm. The following two sections describe empirical research that relates merger activity to either market mispricing or managerial biases. Dong et al. (2006) find that acquirers that are overvalued pay a higher premium for targets, offer stock payments and experience negative announcement day returns. They also find that overvalued targets are more likely to receive stock payments, accept lower premiums and experience lower announcement day returns.

In this study we begin by developing a model that relates the relative optimism of the both the acquiring and target managers to the market valuation of both firms. The explicit recognition that what matters is the relative optimism of both managers is one of this study's contributions. While acquiring manager optimism has been previously demonstrated as an important factor in mergers, we show that the optimism of the target manager is also important and affects both the target manager's preferred form of payment and required premium.

Our model allows us to pinpoint which beliefs affect the preferred method of payment, the premium, or both. For example, we show that the acquiring manager's belief about the target firm's valuation affects the premium the acquirer is willing to pay, but has no effect on the acquirer's preferred form of payment. The latter is affected only by the acquiring manager's beliefs about her own firm's value.[1] On the other hand, the target manager's belief about his own firm's valuation affects the minimum required premium but not the preferred form of payment: a target manager who believes his firm is undervalued will demand a higher premium, but his optimism will not affect his preferred form of payment. The latter is determined only by the target manager's beliefs about whether or not the acquiring firm is fairly valued.

We test the empirical predictions of the model using a sample of 336 merger transactions. The manager's optimism is estimated by their willingness to accept an increase in the risk of their personal portfolio. Specifically, CEO's are labeled as optimistic if they do not exercise options that are significantly in-the-money or they purchase stock in their own company.

The current market valuation is estimated by the price-to-book ratio and also the price to residual-income ratio to include a proxy for possible future growth of the firm. Future growth is estimated from analyst's expectations of future earnings.

Our model offers several additional insights. First, the model is consistent with an environment in which cash mergers are more frequently hostile than stock mergers. In our model, the beliefs that give rise to a preference for cash among acquiring managers at the same time shrink the range of possible premiums acceptable to both the target and acquiring manager. Assuming the likelihood of a friendly merger is increasing in the range of acceptable premiums (combined with our empirical result that acquiring managers appear to dictate the method of payment), then stock mergers are more likely to be friendly.

Most models predict that either cash or stock is optimal, but rarely predict an optimal combination. Our model offers an explanation for why acquiring managers sometimes choose combinations of cash and stock. Because the acquiring manager's beliefs affect her preferred method of payment, the target manager's beliefs about the acquiring firm's valuation will be a function of the acquirer's chosen method of payment. If the acquiring manager can convince the target that the acquiring firm is fairly priced by substituting cash for some of the stock, this reduces the minimum premium the acquiring firm must pay. We show that when the two managers hold different prior beliefs about the acquiring firm's value, the extra cost to the acquirer of substituting cash for stock (when stock is optimal) can be more than offset by the lower premium that the target manager is willing to accept.

The Model

Our model extends the ideas of Shleifer and Vishny (2003) and Malmendier and Tate (2008) by examining mergers when managers are biased and markets are simultaneously misvalued. It centers on the merger of two firms, the Acquirer (*A*) and the Target (*T*), and the future merged combination of these firms (*C*). The perceived value of each firm is indexed by the subscript *i*, which denotes the perspective of either the market (*m*), the acquiring manager (*a*) or the target manager (*t*). In the next section, the subscript *j* is also used to index the relative bias of each manager. If a merger is successful, the value of the combined firm from the viewpoint of *i* will be:

, (1)

where e_i is the synergy gain expected from participant *i*, and *k* is the cash paid by the acquirer to the shareholders of the target firm.[2]

In addition to the cash, the target shareholders can also receive a proportion s_t of the combined firm (acquiring shareholders will receive $(1 – s_t)$ of the new firm). The premium paid by the acquirer, *p*, is defined as the value of the shares plus cash minus the current market value of the target. Therefore,

, (2)

and the proportion of the new firm given to the target shareholders will be

. (3)

It should be noted that the target's share of the combined firm is calculated using the current market values of the acquirer and the target. This assumption is made because share values in merger transactions always use current market values.[3]

The Acquirer's Decision

We begin the analysis by looking at the decision process of the acquiring manager. The manager will desire a merger if the value of the manager's claim after the merger exceeds the value before the merger. The acquiring manager's problem is: subject to , , and . Constraints (i) – (iii) require that the acquiring manager, target manager and target shareholders, respectively, be at least as well off after the merger as they are before. We assume that beliefs vary among individuals and that these belief differences produce different asset valuations.[4]

To capture the acquiring manager's relative beliefs about valuations, we define the acquiring manager's valuation divided by the market valuation as:

— — —

These ratios represent the bias of the acquiring manager, relative to the market, for the value of her own firm (A), the target firm (T) and the synergies that will be obtained by the merged firm (e). Defining the ratios separately for the acquiring firm, target firm and the synergy gains allows us to separately examine belief differences for each component contributing to the merged firm's value. The maximum premium the acquiring manager can pay while satisfying constraint (i) is:

. (4)

Equation (4) establishes an upper bound for the premium the acquiring manager is willing to pay. The first term is standard: the acquiring manager will not pay more for the target firm than the total expected synergy gains. The second term, in square brackets, is affected by the acquiring manager's beliefs about her own firm's value and the synergy gains, while the last term involves her beliefs about the relative valuation of the target. We now consider the implications of specific beliefs.

Case A.1: Unbiased Beliefs (also corresponds to full rationality, efficient markets)

If the acquiring manager and market have identical views on the synergies and the value of each firm, then equation (4) simplifies to . Any decision to merge will be made when the acquiring manager can pay a premium smaller than the synergies that she expects from the merger. This result rests purely on the notion that mergers exist to capture the synergies between two companies. As long as the cost (premium) of the merger is less than the synergies, it is optimal for the acquiring manager to pursue the acquisition. Also in this case, the maximum premium is independent of the method of payment.

Case A.2: Acquiring Manager Believes Own Firm is Overvalued (

If the stock market places a greater value on the acquiring firm than the acquiring manager, this implies that the manager believes her stock is overvalued and will eventually decline towards her perceived value, . The maximum premium the acquiring manager will offer depends on her beliefs about the valuation of the target. First, suppose the acquiring manager believes the target is also overvalued (without loss of generality assume). Then,

. (5)

Equation (5) indicates that the method of payment matters, since the premium is maximized by setting k=0. That is, the acquiring manager can afford to pay higher premium by issuing overvalued stock than by offering cash (whether the target manger also prefers stock is addressed in the next section). The second term captures the "penalty" from using cash instead of overvalued stock.

Next, consider the case where the acquiring manager believes the target firm is fairly priced ( ). In this case,

. (6)

The acquiring manager will still prefer stock to cash, but a third force is now at work: the acquiring manager's ability to transfer some of her own firm's overvaluation to target shareholders enables her to pay a premium in excess of the synergy gains when using stock. The stronger the acquiring manager's belief about her own firm's overvaluation, the larger the premium she can afford to pay. This leads to the following three predictions which are consistent with Shleifer and Vishny (2003):

Prediction 1: *The takeover premium is positively related to the acquiring firm's market valuation and negatively related to the target firm's market valuation.*

It is apparent that a stock merger will be preferred by the acquiring manager in order to transfer some of her own firm's overvaluation to the target shareholders.

Prediction 2: *For a given premium, an acquiring manager will prefer to offer a stock payment when she believes her own firm is overvalued* .

The analysis also provides a basis for the anticipated stock market returns on the day of the merger announcement. Market optimism about the value of the acquiring firm may lead to a premium in excess of synergy gains.

Prediction 3: *Market optimism about the value of the acquiring firm leads to lower announcement day returns for the acquiring firm.*

Case A.3: Acquiring Manager Believes Own Firm is Undervalued

Roll (1986) argues that managerial hubris may help explain why a manager purchases another firm when it is not a value increasing proposition. A manager who suffers from hubris essentially believes that the more she manages the greater the value she can create. In this section we consider an acquiring manager who believes his firm is undervalued. This could be due to optimism, overconfidence, hubris, or some other factor. We are not interested in the causes of such a belief, but the consequences on manager behavior.[5] If the acquiring manager believes the target is fairly valued) then as before

. (7)

In this case, the maximum premium the acquirer will pay is in an all-cash deal. However, if the acquiring manager believes the target is also undervalued, she will be willing to pay more. Again, suppose without loss of generality that , so that

. (8)

The maximum premium is an increasing function of the cash component of the deal and the degree of perceived underpricing. This leads to the following predictions:

Prediction 4: *An acquiring manager who is relatively optimistic about the value of her own firm will prefer cash as the method of payment in a merger.*

Equation (8) shows that optimism on the part of the acquirer is positively related to premium in cash mergers (since the premium is maximized by setting $s = 0$). However, when stock financing is added to the merger, the positive relationship is weakened.

Prediction 5: *An acquiring manager who is relatively optimistic about the value of her own firm will offer a higher premium for mergers financed with cash.*

Prediction 5 qualifies the prediction offered by Malmendier and Tate (2008) that optimistic acquiring managers will offer higher premium regardless of the method of payment. Here, the optimistic acquiring manager recognizes that relative to cash, her own stock is 'undervalued currency'. In fact, holding constant the value of the target firm, the optimistic acquiring manager will offer a *smaller* premium in stock deals than a pessimistic manager who believes her firm is overvalued. Because we expect the acquiring firm's announcement day return to be lower for optimistic managers, especially when cash is used as the method of payment. This leads to the following prediction.

Prediction 6: *An acquiring manager who is optimistic about the value of her own firm will experience a lower announcement day return than an unbiased manager, especially for cash-financed mergers.*

Case A.4: The Optimistic Manager in an Overvalued Market

Because we model the beliefs of the acquiring manager relative to the market, the case where the acquiring manager is relatively optimistic applies equally to all cases in which the acquiring manager is relatively more optimistic than the market (e.g. a rational manager in an undervalued market; an optimistic manager in an efficient market; or even an optimistic manager in an overvalued market, provided the manager is relatively more optimistic than the market). Likewise, the case of a relatively pessimistic acquiring manager applies to all situations where the market is relatively more optimistic than the manager herself. Therefore,

Prediction 7: *An acquiring manager will prefer stock mergers when the market is relatively more optimistic than the acquiring manager about the value of the combined firm.*

In the Shleifer and Vishny (2003) model, target firm undervaluation is a necessary, but not a sufficient, condition for cash deals. Here, cash deals can occur if the target is overvalued provided the acquirer is relatively more optimistic than the market about the degree of overvaluation of the combined firms.

The Target Manager's Decision

We begin the analysis by looking at the decision process of the target manager. We make the simplifying assumption that the target manager negotiates a premium on behalf of the shareholders. We also abstract from the bargaining process and solve for the minimum takeover premium that the target manager will accept. Generally, this is the minimum premium for which there is a perceived increase to the current shareholder wealth, or for which the target shareholders' share of the combined firm plus any cash they receive in the merger is greater than their value before the merger. This occurs when

Further simplification leads to

. (9)

Case B.1: Unbiased Beliefs (also corresponds to full rationality, efficient markets)

If the target manager and market have rational views of the value of each firm then , and the target manager will prefer any positive premium to no merger. This simple result draws from the economic notion of selling an asset whenever one can receive a price above its fundamental value. In this particular specification, the method of payment does not affect this result, and the choice of payment is irrelevant to the post-merger value of the target.

Case B.2: Target Manager Believes Own Firm is Overvalued (

If the target manager believes that his own firm's stock is overvalued, this presents the target manager with an opportunity to capitalize on the misvaluation.[6] We assume that the target manager believes that his stock value will eventually decline toward his perceived fundamental value T_t.

The target manager's beliefs about the valuation of the acquiring firm are important. First, suppose the target manager believes the acquiring firm is overvalued. Then, the minimum premium (assuming without loss of generality) is

—— . (10)

In an all-stock deal (k = 0), , but for an all-cash deal the minimum premium is actually negative. That is, the target manager is willing to sell his firm for less cash than the current market value, because doing so transfers some of his own firm's overpricing to the acquiring shareholders. More generally, the target manager will accept the merger if the loss that he transfers to the acquiring shareholders, plus the premium, exceeds the loss that is obtained from the acquiring firm (via the acquiring firm's overvalued shares).

Prediction 8: *For a given premium, a target manager will prefer a cash payment when he believes that the acquiring firm is overvalued* .

This result contrasts with Prediction 2, which states that the acquiring manager prefers to pay in stock when her firm is overvalued. Therefore, our empirical analysis in the next section tests the effect of market optimism on the method of payment to see whether one manager systematically dictates the form of payment in the merger. Furthermore, this prediction also contrasts with the prediction by Dong et al. (2006) in which they hypothesize that target managers will demand stock payments when the acquirer or the target are overvalued.

On the other hand, if the target manager believes the acquiring firm is fairly priced, then the minimum premium is unaffected by the method of payment. However, the target manager is still willing to accept a negative premium satisfying the condition:

—— . (11)

This analysis also provides a basis for predicting the stock market returns on the day of the merger announcement when the market is optimistic about the value of the target firm. Because the target manager is willing to accept a lower takeover premium when he believes his firm is overvalued, we have:

Prediction 9: *The target firm will experience lower announcement day returns when the target manager believes his own firm is overvalued* .

Case B.3 Target Manager Believes Own Firm Undervalued

We next analyze the case where the target manager believes his firm's fundamental value exceeds its market value. This could be due to optimism, overconfidence, hubris, or some other factor. Again, we are not interested in the causes of such a belief, but its effects on managerial behavior. If the manager believes the acquirer is also undervalued (assume *),* then

—— . (12)

The minimum premium is now positive for cash deals, and all else equal, the target manager demands a higher premium for cash-financed deals than stock deals. This is because (by assumption) he views $1 of the (undervalued) acquiring firm's stock as worth more than $1 of cash.

Prediction 10: *A target manager who is relatively optimistic about the value of the acquiring firm will demand a higher premium for mergers financed with cash than with stock.*

If the target manager believes the acquiring firm is fairly priced, then

$$\text{——} \quad . \tag{13}$$

While the minimum premium is still positive, the target manager is indifferent between stock and cash as the form of payment. By comparing equations (12) and (13) to (10) and (11), we see that relatively optimistic target managers (who believe their firm is undervalued) will demand a higher premium than relatively pessimistic target managers, regardless of what they believe about the relative pricing of the acquiring firm, and regardless of the method of payment. Therefore, we expect the target announcement day return to be higher for optimistic target managers, especially when cash is used as the method of payment.

Prediction 11: *A target manager who is optimistic about the value of his own firm will demand a higher premium and experience a higher announcement day return than an unbiased target manager.*

Case B.4: The Optimistic Manager in an Overvalued Market

Because we model the beliefs of the target manager relative to the market, the case of target manager optimism applies to all cases in which the target manager is relatively more optimistic than the market (e.g. a rational manager in an undervalued market; an optimistic manager in an efficient market; or even an optimistic manager in an overvalued market, provided the manager is relatively more optimistic than the market). Therefore,

Prediction 12: *A target manager who is optimistic about the value the combined firm will prefer cash mergers when the market is relatively more optimistic than the target manager about the value of the combined firm.*

That is, given an optimistic manager, market optimism must exceed the target manager's level in order for cash to be the preferred payment.

Data and Methodology

Merger data is from the Security Data Corporation's (SDC) Global Merger and Acquisition Database. The sample is restricted to US publicly listed firms that have acquired another US publicly listed firm between 1994 and 2004. Both the acquirer and the target must be listed in the ExecuComp database to ensure CEO portfolio data availability. Therefore, the set of mergers is limited to those including firms within the S&P 1500. Firm level data from COMPUSTAT is used to supplement the merger data. The total sample is comprised of 336 mergers with an average deal value of $3,963.7 million. Of those mergers, 147 are stock mergers and 109 are cash mergers. The remaining 80 mergers are a mixture of stock and cash.

Measure of Market Misvaluation

Following Dong et al. (2006), we use two market misvaluation measures: (1) the standard price-to-book value of equity ratio (*P/B*) and (2) the price to residual-income model ratio (*P/RIM*). Both *P/B* and *P/RIM* have been used in numerous studies as predictors of a firm's abnormal future returns.[7] Lee, Myers, and Swaminathan (1999) have shown that *P/RIM* is better at forecasting future stock returns than *P/B*; however, *P/RIM* has the disadvantage of using analyst's estimates in the calculation which may be inherently biased or may be correlated with current market conditions. Thus we use both ratios.

The *P/B* ratio is calculated by dividing the firm's current market value by the book value of the firm's equity at the end of the prior fiscal year. Both values are derived following Fama and French (2002) where market value is calculated by multiplying the stock price one month prior to the merger announcement by the number of common shares outstanding (COMPUSTAT Annual Item 25). Book value of equity is calculated as total assets (Item 6) minus total liabilities (Item 181) and preferred stock (Item 10) plus deferred taxes (Item 35) and convertible debt (Item 79).[8] Finally, the *P/B* ratio is winsorized at the 1% tails; and following Dong et al. (2006), the negative *P/B* ratios are replaced with the maximum *P/B* ratios in the sample. The mean *P/B* ratio for the acquirer sample is 4.34 and the mean for the target sample is 3.59.

The *P/RIM* ratio is derived from the *P/B* ratio with the addition of analysts' estimates to proxy for a firm's future growth potential. Book value follows traditional accounting rules to capture historical cost and does not allow for any measure of future growth. The residual-income model (*RIM*) was created by Ohlson (1995) to combat this problem and capture growth potential. Ohlson begins with book value and adds to it the present value of expected residual income, calculated by first taking the difference between the return on equity and the cost of equity and then multiplying by book value, resulting in:

$$RIM = B\ \ t\ \ + \sum_{i=1}^{\infty} \frac{E_t\ \ ROE\ \ t+i\ \ - r_e\ \ t}{\left[1 + r_e\ \ t\ \right]^i} B\ \ t+i-1\ \ , \tag{14}$$

where *B* is the book value of equity, *ROE* is the return on equity, and r_e is the firm's cost of equity.

Some assumptions are required to replace the infinite sum and compute equation (14). Lee, Myers, and Swaminathan (1999) find that their estimates of *RIM* do not change significantly when their forecast horizon is increased beyond three years. Under this assumption that residual income beyond the third year is constant, equation (14) becomes

$$RIM(t) = B(t) + \frac{f^{ROE}(t+1) - r_e(t)}{1+r_e(t)} B(t) + \frac{f^{ROE}(t+2) - r_e(t)}{\left[1+r_e(t)\right]^2} B(t+1) + TV, \tag{15}$$

where f^{ROE} is the forecasted return on equity and *TV* is the estimated terminal value beyond year 2 calculated as

$$TV = \frac{f^{ROE}(t+3) - r_e(t)}{\left[1+r_e(t)\right]^2 r_e(t)_e} B(t+2). \tag{16}$$

The forecasted return on equity is derived from analysts' estimates of earnings per share (EPS) taken from the I/B/E/S database and constructed as

$$f^{ROE}(t+i) = \frac{f^{EPS}(t+i)}{B(t+i-1)}, \tag{17}$$

where f^{EPS} is the forecasted EPS. When an EPS forecast is not available, the previous forecast is multiplied by the growth rate of earnings which is also supplied by I/B/E/S. The future book values of equity are computed as

$$B(t+i) = B(t+i-1) + f^{EPS}(t+i) - f^{DPS}(t+i), \tag{18}$$

where f^{DPS} is the forecasted dividends per share calculated as

$$f^{DPS}(t+i) = f^{EPS}(t+i) \times k, \tag{19}$$

where *k* is the current dividend payout ratio which is determined by

$$k = \frac{D(t)}{EPS(t)}, \tag{20}$$

and *D(t)* and *EPS(t)* are the current dividend per share and earnings per share, respectively. To compensate for companies with negative EPS, dividends are divided by 6% of total assets to calculate the dividend payout ratio (in accordance with Lee et al., 1999). Finally, the cost of equity is calculated for each firm by the traditional CAPM of:

$$r_e = r_f + \beta_i (r_m - r_f), \tag{21}$$

where r_f is the one-month T-bill rate, r_m is the average annual return on the CRSP value-weighted index over the previous 30 years, and β_i is the firm specific beta calculated using five years of monthly returns.

The mean *P/RIM* ratio for the acquirer sample is 3.06 and the mean for the target sample is 2.70. The averages of our P/B ratio's and P/RIM ratio's are somewhat larger than those found in Dong et al. (2006), which may be due to different sample periods: our sample is from 1994 – 2004 whereas their sample is from 1978 – 2000. The correlation for the acquirer is 30% and the target is 16%. Dong et al. (2006) report correlations of 33% for acquirers and 20% for targets. These low correlations suggest that both measures may offer orthogonal information about market misvaluation.

Measure of Managerial Optimism

Following the approach implemented by Malmendier and Tate (2005), we proxy managerial optimism by observing CEO personal investment in their own company. Data on CEO portfolios is obtained from the ExecuComp database which includes information on option and stock holdings at the fiscal year end for companies in the S&P 1500. CEOs who fail to reduce their exposure to company specific risk are classified as optimistic. This exposure to company risk is then exacerbated by the CEO's large human capital investment in the firm resulting in under-diversified portfolios. Thus, when CEOs fail to take advantage of opportunities to reduce their exposure to firm specific risk or when they increase their exposure, they are classified as optimistic.

The first measure of optimism is proxied by a CEO who holds vested options that are significantly in-the-money. Hall and Murphy (2002) derive the conditions under which it is optimal for unbiased CEOs to exercise their options given their individual wealth and degree of risk-aversion. Malmendier and Tate (2005) utilize this optimality and define as optimistic any CEO who does not exercise stock options that are at least 67% in-the-money in any year after the options have vested. They further control for chance instances of nonexercise by requiring that the CEOs fail to exercise in at least two years in which they had options that were exercisable and significantly in-the-money.

To calculate the percentage that the CEO's vested options are in-the-money we divide the profit from exercising their options (fiscal year-end stock price minus the average exercise price of their options) by the average exercise price of the CEO's in-the-money options. The average exercise price of the options is calculated in accordance with Core and Guay (2002) by taking the value of the CEO's vested options and dividing by the number of vested options and then subtracting this value from the fiscal year-end stock price. They note that their methodology may introduce a downward bias when computing the exercise price when stock prices are decreasing because they assume that out-of-the-money options have exercise prices equal to the current fiscal year-end stock price. To correct for this bias, we modify the methodology of Malmendier and Tate (2005) by classifying CEO's as optimistic when they fail to exercise, at least twice, options that are 100% or more in-the-money. Using this measure we find 61% of the acquirers are classified as optimistic (compared with 51% in the sample of Malmendier and Tate 2005) and 45% of the target managers are classified as optimistic. Each manager's optimism is negatively correlated with their respective market valuations.[9]

The second optimism measure is derived from the Hall and Murphy (2002) framework and follows from their assertion that under-diversified CEOs should not increase their exposure to company risk by purchasing company stock. Therefore, if a CEO bought company stock on net in more years than she/he sold stock on net over her/his tenure, they will be classified as optimistic. The data for CEO stock holdings is obtained from ExecuComp and any shares obtained through stock grants are excluded.

Under the 'Stock Purchaser' measure, we find 56% of the acquirers are classified as optimistic (compared with 61% in the Malmendier and Tate (2005) sample) and 47% of the target managers are classified as optimistic. Optimistic acquiring managers have little correlation with our market measures and the optimism of target managers also exhibit low correlation with the market. The correlation between both managers is 0.088. The 'Option Holder' measure and the 'Stock Purchaser' measure have a correlation for the acquirer and target of 0.159 and 0.241, respectively.

We include several control variables in our regressions to isolate the effects that optimism has on merger characteristics. These control variables are similar to those in Dong et al. (2006) who test investor misvaluation on mergers. Because synergies among related firms can be a major motive in mergers, we include a dummy variable that is equal to one when a merger is conducted among firms in the same industry. We define industry based on 3-digit SIC code. In addition, acquirer industry dummy variables are included to control for differing effects across industries. Year dummy variables are also included to control for changes in merger characteristics across time.

We use the log of both acquirer size and target size, measured as market value, as control variables. The average acquirer market value is $55,287 million and the average target market value is $8,782 million. The final control variable is the leverage of the acquirer defined as the total debt of the acquirer divided by the total assets of the acquirer. The leverage variable helps control for the financing constraints that could influence the behavior of the acquirer as explained by the Jensen (1986) free cash flow argument. The mean leverage for our sample is 0.180.

Empirical Results

This section examines the impact of managerial optimism and market misevaluation on (1) the method of payment utilized in mergers; (2) the size of the premium; and (3) the market reaction to the merger announcement.

Method of Payment

According to our model, acquiring managers will prefer to use their stock as payment when they believe their firm is overvalued (Prediction 2). In contrast, target managers will prefer to receive cash when they believe the acquirer is overvalued (Prediction 8). Therefore, tests of the method of payment used in different market conditions will determine whether one manager predominantly decides on the method of payment in mergers.

Table 1 provides the results for logistic regressions of our proxies for market valuation on stock and cash mergers. All of the coefficients of price-to-book (*P/B*) and price to residual-income (*P/RIM*) for the regression on stock mergers are positive. Of the *P/B* coefficients, the acquirer's stock valuation is significant at the 1% level and the *P/RIM* results are also significant for the acquirer. The cash merger results show negative and significant coefficients. The acquirer's valuation measures are significant at the 1% and 5% level for *P/B* and *P/RIM*, respectively. These results indicate that stock is used in a relatively optimistic market and cash in a pessimistic one. Because these patterns correspond to the acquiring manager's preferred method of payment, acquiring managers have primary influence on the method of payment used in mergers.

Table 1. Logistic Regressions of Market Optimism on Cash & Stock Mergers (n = 336)

	(1)	(2)	(1)	(2)
Acquirer *P/B*	0.150***		-0.133***	
	(4.830)		(-3.922)	
Target *P/B*	0.051**		-0.038*	
	(2.326)		(-1.698)	
Acquirer *P/RIM*		0.059***		-0.066**
		(2.642)		(-2.401)
Target *P/RIM*		0.037		-0.044*
		(1.540)		(-1.688)
Non-diversifying	0.300	0.244	-0.332*	-0.298
	(1.603)	(1.332)	(-1.713)	(-1.560)
Log(Acquirer Size)	-0.270***	-0.211***	0.267***	0.238***
	(-3.584)	(-2.934)	(3.449)	(3.127)
Log(Target Size)	0.261***	0.289***	-0.265***	-0.315***
	(3.386)	(3.772)	(-3.233)	(-3.736)
Leverage	-0.802	-0.701	1.190	0.990
	(-1.011)	(-0.948)	(1.469)	(1.281)
McFadden R2	0.3242	0.2831	0.3281	0.3054

z-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

Table 2 tests the prediction that optimistic acquiring managers will be more likely to offer cash in mergers (Prediction 4). The table shows that optimism does increase the acquiring manager's propensity to use cash. These results indicate that optimism alone does not have a significant influence on the form of payment in a merger.

While market valuation affects the choice of payment more than relative managerial optimism, our model also predicts that there will be an interaction between managerial optimism and market valuation. Because managerial optimism is measured on a binary scale, we cannot test Predictions 7 and 12 by calculating the difference between managerial and market optimism to determine which is relatively more optimistic. However, relative optimism does suggest that market valuation will be positively related to optimism of the acquirer manager in stock mergers and positively related to optimism of the target manager in cash mergers. To perform a test of this relation we use the following logistic regression:

$$\Pr\left(Y=1|O,M,X\right)=G\left(\beta_1+\beta_2 O+\beta_3 M+\beta_4\left(O\cdot M\right)+X'B\right) \quad (22)$$

where O is the optimism measure of either the acquirer or the target (Option Holder or Stock Purchaser), M is the combined market valuation measures (*P/B* or *P/RIM*), and X is the set of control variables. Y is a binary variable equal to 1 if the merger was completed with either all cash or all stock where noted. We assume that G follows the logistic distribution. The null hypothesis is that β_4, the coefficient on the interaction of market valuation and CEO optimism, is equal to zero.

Table 2. Logistic Regressions of Acquirer Optimism on Method of Payment (n = 336)

	(1)	(2)
Option Holder	0.127	
	(0.665)	
Stock Purchaser		0.097
		(0.508)
Non-diversifying	-0.334*	-0.336*
	(-1.783)	(-1.790)
Log(Acquirer Size)	0.171**	0.153**
	(2.409)	(2.084)
Log(Target Size)	-0.259***	-0.244***
	(-3.308)	(-3.123)
Leverage	0.978	1.003
	(1.294)	(1.335)
McFadden R2	0.2770	0.2766

z-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

To compute the regression given in Equation (22) we need a measure for the combined market value of the acquirer and the target. We, instead, form a modified *P/B* and *P/RIM* for the combination of the acquirer and target. Therefore, the combined market value M will be calculated as

$$M = \frac{P_a + P_t}{B_a + B_t} \text{ or } \frac{P_a + P_t}{RIM_a + RIM_t}, \tag{23}$$

where P_i is market value, B_i is book value, RIM_i is the residual income model calculation given in Equation 14, and i represents the acquirer (a) or the target (t).

The results for the effects on a stock merger for acquiring managers are shown in Table 3. The coefficients on the level of managerial optimism are negative and insignificant. The coefficients on market valuation are positive and significant (with the exception of regression (4) which has a significance level of 12%). The interaction term between market valuation and managerial optimism is significant in three of the four models. This is consistent with the conjecture that optimistic acquiring managers will prefer stock mergers when the market is sufficiently overvalued (i.e. when the market is relatively more optimistic than the acquiring manager about the value of the combined firm).

Table 4 tests whether target manager optimism influences the likelihood of a cash merger. The coefficients of the level of optimism of the target manager are negative and insignificant (with the exception of the second regression where it is positive and insignificant) implying a weak penchant for stock mergers. The coefficients on market valuation are negative and significant. The interaction term between the market and managerial optimism is significant in two of the four models at the 10% level providing some support for a positive relationship between optimistic target managers and market valuation in cash mergers.

The relationship between managerial optimism and market valuation could give credence to the theory that optimistic manager's view of the market is skewed by their own perceptions. Acquiring managers prefer offering stock when the market is overvalued; however, the degree of overvaluation that they perceive in the marketplace is affected by their level of optimism. In other words, we conclude that optimistic acquiring managers require a higher level of relative market valuation before offering stock in mergers. Conversely, optimistic target managers gain by accepting cash in an overvalued market. It also appears target managerial optimism alone does not influence the method of payment in mergers, however, it does seem to influence the level at which the manager views stock market misvaluation.

Table 3. Logistic Regressions of Acquirer & Market Optimism on Stock Mergers (n = 336)

	(1)	(2)	(3)	(4)
Option Holder	-0.219	-0.440		
	(-0.631)	(-1.447)		
Stock Purchaser			-0.330	-0.418
			(-0.950)	(-1.473)
Combined (Acquirer & Target) *P/B*	0.194***		0.188***	
	(3.845)		(2.899)	
Combined (Acquirer & Target) *P/RIM*		0.086**		0.061
		(2.474)		(1.626)
(CEO Measure) x (Market Valuation)	0.133*	0.193***	0.055	0.110**
	(1.719)	(2.766)	(0.702)	(1.985)
Non-diversifying	0.305	0.241	0.278	0.247
	(1.609)	(1.293)	(1.488)	(1.341)
Log(Acquirer Size)	-0.334***	-0.302***	-0.292***	-0.287***
	(-4.181)	(-3.816)	(-3.609)	(-3.483)
Log(Target Size)	0.309***	0.317***	0.272***	0.315***
	(3.868)	(3.933)	(3.433)	(3.915)
Leverage	-0.825	-0.911	-0.709	-0.529
	(-1.042)	(-1.199)	(-0.912)	(-0.708)
McFadden R2	0.3335	0.3025	0.3151	0.2910

z-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

Table 4. Logistic Regressions of Target & Market Optimism on Cash Mergers (n = 336)

	(1)	(2)	(3)	(4)
Option Holder	-0.487	0.012		
	(1.443)	(0.045)		
Stock Purchaser			-0.505	-0.104
			(-1.595)	(-0.388)
Combined (Acquirer & Target) *P/B*	-0.246***		-0.275***	
	(-4.178)		(-4.237)	
Combined (Acquirer & Target) *P/RIM*		-0.087**		-0.085**
		(-2.288)		(-2.138)
(CEO Measure) x (Market Valuation)	0.153*	0.028	0.136*	0.012
	(1.850)	(0.500)	(1.875)	(0.234)
Non-diversifying	-0.267	-0.298	-0.270	-0.302
	(-1.366)	(-1.570)	(-1.392)	(-1.579)
Log(Acquirer Size)	0.273***	0.233***	0.325***	0.247***
	(3.208)	(2.878)	(3.938)	(3.136)
Log(Target Size)	-0.301***	-0.301***	-0.315***	-0.304***
	(-3.627)	(-3.569)	(-3.769)	(-3.622)
Leverage	0.976	0.926	0.879	0.963
	(1.214)	(1.212)	(1.098)	(1.256)
McFadden R2	0.3377	0.2941	0.3377	0.2929

z-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

Premium

We now consider factors that influence the size of the premium paid to target shareholders. We measure premium as the price per share offered by the acquirer divided by the target's stock price five days prior to the announcement date of the merger. The average premium for our sample is 29.063% which is consistent with other studies (e.g. Dong et al. 2006).

Our model suggests that the premium is affected by the degree of misvaluation in the market. Table 5 tests the prediction that overvaluation of the acquiring (target) firm leads to a higher (lower) premium as stated in Prediction 1. The coefficients on the market valuation of the acquirer are 0.332 for *P/B* and 0.265 for *P/RIM*. While they are both positive, they are insignificant. Target market valuation coefficients of -1.165 for *P/B* and -1.108 for *P/RIM* are both significant at the 1% level. These results are similar to those found by Dong et al. (2006) and consistent with Prediction 1.

Table 5. Least Squares Regression of Market Optimism on Premium (n = 336)

	(1)	(2)
Acquirer *P/B*	0.332	
	(0.861)	
Target *P/B*	-1.165***	
	(-3.735)	
Acquirer *P/RIM*		0.265
		(1.392)
Target *P/RIM*		-1.108***
		(-3.220)
Non-diversifying	1.603	1.493
	(0.600)	(0.555)
Log(Acquirer Size)	-0.290	-0.003
	(-0.272)	(-0.003)
Log(Target Size)	1.046	0.455
	(0.946)	(0.410)
Leverage	-25.007**	-22.330**
	(-2.331)	(-2.079)
Adjusted R2	0.1478	0.1392

t-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

Table 6 tests whether the premium is affected by managerial optimism. Prediction 5 states that an optimistic acquiring manager will pay a larger premium when paying cash and Prediction 10 states that a target manager who is optimistic about the value of the acquiring firm will demand a larger premium in cash mergers. To test these predictions we use the following general regression:

$$\text{Premium} = G\left(\beta_1 + \beta_2 O + \beta_3 Cash + \beta_4 \left(O \cdot Cash\right) + X'B\right) \quad (24)$$

where O is the optimism measure of either the acquirer or the target (Option Holder and Stock Purchaser), *Cash* indicates that cash was used to finance the merger, and X is the set of control variables. The null hypothesis is that β_4, the coefficient on the interaction of cash and CEO optimism is equal to zero.

The coefficients on managerial optimism are positive but insignificant. The coefficients on cash are also both insignificantly different from zero. The interaction coefficients are 13.683 and 9.935 with significance at the 5% level and 10% level, respectively. The significance of the interaction term indicates that optimistic acquiring managers offer larger premiums in cash mergers than stock mergers. These results are noteworthy because they qualify the evidence in Malmendier and Tate (2008) that acquiring manager optimism leads to higher premiums. While managerial optimism is significant when tested alone, it becomes insignificant when the interaction with cash is included.

Table 6. Least Squares Regressions of Premium on Managerial Optimism & Cash Merger (n = 336)

	Acquirer		Target	
	(1)	(2)	(3)	(4)
Option Holder	0.587		2.467	
	(0.182)		(0.789)	
Stock Purchaser		0.834		3.252
		(0.265)		(1.041)
Cash Merger	-2.268	0.330	0.055	3.138
	(-0.510)	(0.076)	(0.015)	(0.862)
(CEO Measure) x (Cash Merger)	13.683**	9.935*	12.489**	7.480
	(2.438)	(1.775)	(2.265)	(1.380)
Non-diversifying	0.935	1.459	2.140	1.709
	(0.346)	(0.537)	(0.795)	(0.629)
Log(Acquirer Size)	0.044	-0.798	-0.626	-0.498
	(0.041)	(-0.730)	(-0.601)	(-0.474)
Log(Target Size)	0.718	1.379	0.910	1.049
	(0.636)	(1.212)	(0.817)	(0.930)
Leverage	-23.303**	-22.996**	-23.964**	-21.996**
	(-2.164)	(-2.136)	(-2.239)	(-2.057)
Adjusted R2	0.1392	0.1270	0.1481	0.1320

t-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

Table 6 also provides the results for the effects of target manager optimism and a cash payment on the amount of premium paid. The coefficients on target managerial optimism are positive but are insignificant.[10] The variables on cash are both positive and insignificant from zero. The interaction coefficients are 12.489 and 7.480 with the first one significant at the 5% level. These results suggest that optimistic target managers seek larger premiums in cash mergers.

Market Reaction

The reaction of the market to merger announcements can be used as a gauge to measure the market's perception of the merger. We estimate the market's reaction to the merger by calculating the cumulative abnormal return (CAR) around the announcement date of the merger. The CAR is computed using standard event study methodology of Brown and Warner (1985). We employ a three-day event window around the announcement date (day 0). Expected returns are computed assuming that α = 0 and β = 1 in the standard CAPM framework following Fuller et al. (2002). Therefore, we proxy expected returns by the CRSP value-weighted market return. Finally, abnormal returns are calculated to be

$$AR_{it} = r_{it} - r_{mt} \tag{25}$$

where r_{it} is the return for firm i on date t, and r_{mt} is the market return. Then the cumulative abnormal return is computed as

$$CAR_i = \sum_{t=-1}^{1} AR_{it}\ . \tag{26}$$

The acquirer has an average CAR of -2.187% and the target's CAR is 19.105%. These returns are similar to those found in other studies (e.g. Kaplan and Weisbach, 1992) and consistent with the acquirer losing value or breaking even and the target gaining value in mergers.

Table 7 provides the results for regressions of the acquiring firm's CAR on it market value and manager's optimism. The coefficients for the CEO optimism measure are positive but statistically insignificant. This result differs from Malmendier and Tate (2008) who report a significantly negative coefficient. Market optimism exhibits a statistically significant negative relationship with the acquirer CAR having coefficients of -0.235 for *P/B* and -0.118 for *P/RIM*; both are significant at the 5% level. As might be expected, market misvaluation leads to lower announcement day returns providing evidence that the announcement may signal the acquiring manager's belief that her own firm is overvalued. It does not appear, however, that the market is able to perceive the relative optimism of the acquirer. This could be due to the lack of transparency in the marketplace with respect to the manager's optimism level. Alternatively, the market may "buy into" arguments by the acquiring manager that synergy gains are substantial.

Table 7. Least Squares Regressions of Acquirer CAR on Acquirer & Market Optimism (n = 336)

	(1)	(2)	(3)	(4)
Option Holder	0.549			
	(0.552)			
Stock Purchaser		0.299		
		(0.312)		
Cash Merger	1.638	0.888		
	(1.197)	(0.677)		
(CEO Measure) x (Cash Merger)	-0.660	0.591		
	(-0.383)	(0.347)		
Acquirer *P/B*			-0.235**	
			(-1.998)	
Acquirer *P/RIM*				-0.118**
				(-2.070)
Non-diversifying	0.213	0.178	-0.048	0.112
	(0.257)	(0.216)	(-0.060)	(0.138)
Log(Acquirer Size)	-0.044	-0.078	0.171	0.092
	(-0.134)	(-0.232)	(0.523)	(0.285)
Log(Target Size)	-0.527	-0.484	-0.609*	-0.641*
	(-1.523)	(-1.397)	(-1.826)	(-1.921)
Leverage	-2.503	-2.546	-1.019	-2.103
	(-0.756)	(-0.774)	(-0.314)	(-0.649)
Adjusted R2	0.0964	0.0970	0.1073	0.1082

t-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

Table 8 estimates the same basic equation for the target firm. Just as with the acquirer, the overvaluation of the target's stock leads to lower announcement day returns. The coefficient of *P/B* in the regression on Target CAR is -0.921 and is significant at the 1% level; *P/RIM* has a coefficient of -0.644 which is significant at the 10% level. The reaction of the market to both measures of target manager optimism is significant at the 10% level with Option Holder having a coefficient of 4.444 and Stock Purchaser's coefficient equal to 4.687. These results indicate that target shareholders get rewarded in a merger if their CEO is relatively optimistic, and that this reward comes in the form of the increased premium that the optimistic target manager demands (see Table 6).

Conclusion

The empirical results suggest that market valuation of the acquiring firm has a strong influence on the form of payment. The more overvalued the acquiring firm, the more likely that the merger will be financed with cash. The form of payment is not directly affected by the optimism of either manager. However, an optimistic acquiring manager will offer stock if her firm is sufficiently overvalued. The size of the premium is affected by the valuation of the target firm and the optimism of the target manager. Overvalued targets receive lower premiums, while but the premium is increasing with the optimism of the target manager. For cash deals, the premium is also higher when an optimistic acquiring manager is involved. In addition, the observed form of payment corresponds to the acquiring manager's preferences, suggesting that the acquiring manager may dictate the method of payment.

Table 8. Least Squares Regressions of Target CAR on Target & Market Optimism (n = 336)

	(1)	(2)	(3)	(4)
Option Holder	4.444*			
	(1.856)			
Stock Purchaser		4.687*		
		(1.938)		
Target *P/B*			-0.921***	
			(-3.170)	
Target *P/RIM*				-0.644*
				(-1.781)
Non-diversifying	-1.482	-1.718	-1.624	-1.097
	(-0.583)	(-0.675)	(-0.646)	(-0.430)
Log(Acquirer Size)	1.676*	1.787*	1.842*	1.897*
	(1.701)	(1.816)	(1.892)	(1.919)
Log(Target Size)	-1.115	-1.176	-0.687	-1.205
	(-1.058)	(-1.114)	(-0.658)	(-1.137)
Leverage	-10.825	-11.403	-14.763	-12.730
	(-1.078)	(-1.137)	(-1.478)	(-1.262)
Adjusted R2	0.0778	0.0789	0.0992	0.0769

t-statistics are in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%

Notes

1. Hereafter we use the pronoun ‘she’ to refer to the acquiring firm manager and ‘he’ to refer to the target firm manager.
2. Cash is subtracted from the value of the combined firm because the cash departs the firm after the merger. Therefore, in an all cash merger, the combined firm will be equal to the value of the acquirer plus the accompanied synergies (cash would be equal to the target firm). If the acquirer pays a premium equal to the synergies, the value of the combined firm will not have changed from the pre-merger value (i.e. no value is added).
3. Cai and Vijh (2007) have shown that merger premiums are affected by the liquidity of managerial stock and option holdings. While acknowledging the link between managerial compensation and merger premiums, our model abstracts from this effect to focus exclusively on managerial beliefs and market misevaluation.
4. Our model does not attempt to identify the cause of the heterogeneous beliefs or misvaluation. We instead refer to prior literature of how stock market prices may deviate from fundamental values (e.g. Shleifer, 2000) or the tendency of CEOs to overestimate their own skill level and consequently, the value of their company (e.g. Larwood and Whittaker, 1977 and Camerer and Lovallo, 1999).
5. The results of this scenario are the same for an unbiased manager in a pessimistic market.
6. This scenario applies equally to a rational manager in an overvalued market, or a pessimistic manager in a fairly valued market. The key feature is that the manager believes his firm is worth *less* than the market value.
7. See for example, Daniel, Hirshleifer, and Teoh (2003) and Barberis and Huang (2001) for P/B studies and Ali, Hwang, and Trombley (2003) and Frankel and Lee (1998) for P/RIM studies.
8. Redemption value of preferred stock (Item 56) is substituted when preferred stock is missing.
9. The negative correlation of the ‘Option Holder’ measure may imply that managers may utilize some inside information when deciding to exercise their options. However, Malmendier and Tate (2005b) state that this optimism variable is “a reflection of the CEO’s attitude” and not an indicator of her/his belief in the future direction of the stock. They support this assertion by showing that CEOs would have improved their economic position by exercising their options early and investing the proceeds in the S&P 500. Therefore, this measure of CEO optimism is designed to capture the attitude of the CEO and not be associated with market movements.
10. As with the acquiring manager, optimism of the target manager does lead to higher premiums when tested alone but not with the inclusion of the interaction with cash payments.

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Personal Income Tax Evasion Determinants Revisited: An Exploratory Study Using Newly Available Data

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Abstract

In 2010, the IRS released a series on tax evasion running through 2005. Using these data, the *most current available* from the IRS, this study investigates the impact on tax evasion of income tax rates, IRS audit rates, the unemployment rate, the public's approval rating of the President, the Tax Reform Act of 1986, and two variables previously unstudied in a time-series context, namely, the percentage of filed tax returns having itemized deductions and the interest rate on three year Treasury notes. All of these factors are found to exercise statistically significant influences over personal income tax evasion over the study period 1976-2005.

Introduction

It has been known for some time that U.S. households admit to holding only a small fraction, about 15 percent, of the supply of U.S. currency held outside of financial institutions (Cagan, 1958; Bawley, 1982; Carson, 1984; Pyle, 1989; Feige, 1994; Cebula, 2001). It is also known that that some portion of this unaccounted-for currency is held overseas whereas some sizable portion is held domestically and used for transactions involving incomes unreported to the IRS (Feige, 1994; Cebula, 2001; Ledbetter, 2004). This activity is known as the "underground economy hypothesis" (Feige, 1994). The latter hypothesis has in fact led to the evolution of a body of literature addressing income tax evasion. Income tax evasion effectively consists of taxable income that is either unreported or underreported to the IRS, although it also can consist of spurious or inflated tax deductions.

Studies of income tax evasion behavior essentially fall into three categories. First, there are the principally theoretical Models of tax evasion behavior, such as Allingham and Sandmo (1972), Falkinger (1988), Klepper, Nagin, and Spurr (1991), Das-Gupta (1994), Pestieau, Possen, and Slutsky (1994), Caballe and Panades (1997), and Gahramanov (2009). Such studies are often elegant mathematically and in many cases identify variables that *theoretically may* affect tax evasion. However, such studies tend to provide limited guidance regarding the expected magnitudes of the effects of variables they Model.

Second, there are a number of studies that either (a) use questionnaires or (b) undertake experiments, such as Spicer and Lundstedt (1976), Spicer and Thomas (1982), Baldry (1987), Thurman (1991), and Alm, McClelland, and Schulze (1999). These studies are of course empirical in nature, deriving the data largely (if not entirely) from the experiments. Certain of these studies indicate an aversion to the prospect of being audited while others reveal a lack of such risk-averse behavior; still others imply that taxpayers may be averse to tax evasion on moral grounds. Additionally, the incentive provided by higher marginal income tax rates to evade taxation by underreporting income is also revealed in various such studies.

Third, there are those studies that largely or in some cases exclusively adopt what is referred to as "official data," data obtained from the IRS (or its counterpart outside of the U.S.) and/or some other "official," i.e., "government" source. Among the types of information thusly obtained and analyzed are data on tax evasion, tax rates, and audit rates. Such studies endeavor typically to estimate either the size of the "underground economy" or the aggregate degree of tax evasion or to identify the determinants of same (Tanzi, 1982, 1983; Clotfelter, 1983; Carson, 1984; Long and Gwartney, 1987; Pyle, 1989; Feinstein, 1991; Erard and Feinstein, 1994; Feige, 1994; Cebula, 2001, 2004, 2008; Ali, Cecil, and Knoblett, 2001; Ledbetter, 2004; Connelly, 2004; Christie and Holzner, 2006; Alm and Yunus, 2009; Cebula and Coombs, 2009).

Within the latter context, this exploratory study seeks to add to the rich literature on income tax evasion by empirically investigating determinants of aggregate federal personal income tax evasion in the U.S. using the *most current data available* from the IRS. To date, the empirical literature has effectively failed to investigate determinants of aggregate income tax evasion in the U.S. for recent years; indeed, except for a *single*, somewhat narrowly focused study that investigated tax evasion and government-spending-induced budget deficits through the year 2001 (Cebula and Coombs, 2009), the most recent year considered in the tax-evasion-determinants literature is in fact 1997 (Cebula, 2008; Ali, Cecil, and Knoblett, 2001; Alm and Yunus, 2009). However, the IRS (2010) has very recently released new *time-series* data on tax evasion running through the year 2005. Using these and other new data derived by the IRS (2009A; 2009B; 2010), the present study seeks to identify key personal income tax evasion determinants through the year 2005.

In addition to investigating the effects of the most commonly recognized factors that allegedly influence personal income tax evasion, such as income tax rates and IRS audit rates, this study also investigates the potential impacts of the unemployment rate, the public's job approval rating of the President, and the Tax Reform Act of 1986. Furthermore, this project examines the potential tax evasion impacts of two variables previously unstudied in a purely *time-series* context, namely, the percentage of filed tax returns that include itemized deductions and the interest rate yield on three year U.S Treasury notes. Interestingly, all of these factors are found to be statistically significant influences over the aggregate degree of federal personal income taxation over the 30 year period from 1976 through 2005, the most recent several years of which have not been previously investigated.

The framework for the empirical analysis is presented in the next section of this study. The formal empirical analysis is provided in the subsequent section of the study. Finally, the closing section provides an overview of the study findings.

Framework for the Analysis

In this study, the *relative* probability that the *representative* economic agent will *not* report his/her taxable income to the IRS is treated as positively impacted by (an increasing function of) the expected gross benefits to the agent of not reporting income, *eb*, and as negatively impacted by (a decreasing function of) the expected gross costs to the agent of not reporting income, *ec*. Thus, the ratio of the probability of not reporting income to the IRS, *pnr*, to the probability of reporting income to the IRS, (1-*pnr*), is described for the representative economic agent by:

$$pnr/(1-pnr) = f(eb, ec),\ f_{eb} > 0,\ f_{ec} < 0. \tag{1}$$

Expressing probabilities in *relative* terms such as shown in equation (1) possesses the virtue that it thereby reflects the form of the available tax evasion data, i.e., data where (as described below) the aggregate degree of federal personal income tax evasion is expressed in relative terms (IRS, 2010).

As already observed, the gross expected benefits from *not* reporting income to the IRS are hypothesized to be directly related to the federal personal income tax rate (Cagan, 1958; Bawley, 1982; Tanzi, 1982; Clotfelter, 1983; Pyle, 1989; Feige, 1994). To reflect the federal personal income tax rate, most previous studies using official data for the U.S. have adopted either of two alternative measures: an average effective federal personal income tax rate (*ATR*) or the maximum marginal federal personal income tax rate (*MAXT*). In this study, the *MAXT* measure of the income tax rate is adopted because, as argued in Feige (1994), this tax rate is likely to be a more representative measure of the overall tax burden of the personal income tax rate than *MAXT* would be. Accordingly, it is hypothesized, *ceteris paribus*, that:

$$eb = g(ATR),\ g_{ATR} > 0. \tag{2}$$

The Tax Reform Act of 1986 [*TRA*] may have been perceived by at least some portion of the general public as an honest, good faith effort to reform, i.e., to simplify and increase the equity of the Internal Revenue Code. As Musgrave observed (1987, p. 59), "The Tax Reform Act of 1986 is the most sweeping reform since the early 1940s…" Indeed, the *TRA* did introduce a number of reforms, many of which are outlined in broad terms in Ott and Vegari (2003), Barth (1991), and Sanger, Sirmans, and Turnbull (1990). For example, as observed in Ott and Vegari (2003, p. 279), "The Act introduced major cuts in the personal tax rate. When fully effective, only two tax brackets, set at 15 and 28 percent, were to replace the 14 bracket tax schedule with rates in the range of 11 to 50 percent...[while it] broadened the tax base by reducing the itemized deduction." Musgrave (1987, p. 59) further observes that prior to the *TRA*, a slow erosion of the income tax base had been occurring. Musgrave (1987, p. 57) was particularly dismayed by the widening of tax loopholes and the emergence of high income tax shelters that had "…gained momentum in recent years and undermined the public's faith in the income tax." In this vein, Barth (1991) and Sanger, Sirmans, and Turnbull (1990) describe how the *TRA* decreased depreciation benefits from financial investments in residential as well as commercial real estate, established limitations on the tax deductibility of losses from "passive" investments that affected limited partnerships syndications (including those involving real estate ventures), and terminated favorable capital gains treatment of real estate. Musgrave (1987, p. 59) also expressed concern that the "…compounding of the investment tax credit and accelerated depreciation diluted and distorted the base of the corporate income tax." Musgrave (1987, p. 59) asserted that the *TRA* "…reversed these trends, a major accomplishment that all reformers will welcome." Based on Musgrave's (1987) arguments, then, it is expected in the present study that taxpayers might well have favorably regarded the *TRA* and been less resentful of the Internal Revenue Code than before, at least initially. Thus, it is hypothesized here that during the time frame when the *TRA* was enacted and became fully effective (1986-1987) and also received the greatest publicity, reduced taxpayer resentment of the federal income tax system/Internal Revenue Code would/could, at least temporarily, have resulted in a reduced degree of tax evasion, *ceteris paribus*. The reason this reaction to the *TRA* might be only *temporary* is revealed in the words of Slemrod (1992, p. 45), who argues that it would take at least some time for taxpayers "…to learn about and adjust to the new law [the *TRA*]." Consequently, it is hypothesized here that, for the period 1986-1987, the aggregate degree of federal personal income tax evasion was reduced. Accordingly, (2) above is replaced by (3):

$$eb = j(ATR, TRA),\ j_{ATR} > 0,\ \ j_{TRA} < 0. \tag{3}$$

Next, this study seeks to inquire further into an issue studied by Alm and Yunus (2009), who examined a cross-state panel of *individual* income tax returns for the period 1979-1997. In spirit following Alm and Yunus (2009) but dealing with *aggregate time-series* data rather than a panel of individual tax returns, the present study empirically investigates the impact of itemizing deductions on income tax returns on the propensity to engage in income tax evasion. Arguing that the presence of itemized tax deductions would make any individual tax return more complicated for the IRS to investigate or challenge and hence arguing that such itemized deductions created a potential *opportunity* for *individuals* to engage in increased tax evasion, Alm and Yunus (2009) found modest empirical evidence that this factor potentially raised the degree of income tax evasion. This factor has not previously been considered in the related empirical *time-series* literature. Consequently, in order to help fill this potentially important omission, in the present study, it is also hypothesized that the greater the percentage of personal income tax returns *in the aggregate* that includes itemized tax deductions on Form 1040, Schedule A [*PCTITEM*], the greater the expected benefits from itemizing deductions and hence the greater the degree of personal income tax evasion, *ceteris paribus*. Based on this expectation, (3) is replaced by (4):

$$eb = j(ATR, TRA, PCTITEM),\ j_{ATR} > 0,\ j_{TRA} < 0,\ j_{PCTITEM} > 0. \quad (4)$$

Next, based on Alm and Yunus (2009), Gahramanov (2009), and Cebula and Coombs (2009), it is expected that the higher the unemployment rate [*UN*], the greater the expected benefits of personal income tax evasion, *ceteris paribus*. This is based on the reasoning that the higher the unemployment rate, the greater the extent to which the unemployed work in the "underground economy" and hence do not report income. Furthermore, this effect may be reinforced to the extent that a higher unemployment rate creates an incentive even for still-employed people to avoid taxes to the degree that they try to covet extra funds (by under-reporting income) in *anticipation* of a *possible future* lay-off (Alm and Yunus, 2009; Gahramanov, 2009; Cebula and Coombs, 2009). As a result, equation (4) is expanded to equation (5):

$$eb = j(ATR, TRA, PCTITEM, UN),\ j_{ATR} > 0,\ j_{TRA} < 0,\ j_{PCTITEM} > 0,\ j_{UN} > 0. \quad (5)$$

Additionally, there is the issue of the public's job approval rating of the President [*APPROV*]. Following the study of the period *prior to 1998* by Cebula (2008), it is argued here that the higher the public's job approval rating of the President's performance in office, the greater the degree to which there is satisfaction with the President's actions and policies. The latter can be interpreted, at least to some degree, as implying less public resentment towards or greater approval of his various spending and/or tax policies (as well, perhaps, as his other policies). Similarly, the lower the public's job approval rating of the President, the greater the degree to which the public is likely to be dissatisfied with the President's actions and policies. In turn, it can be reasonably argued that the latter can be interpreted, to at least some extent, as implying greater resentment of or less public support of his various spending and/or tax policies (as well, perhaps, as his other policies). Stated somewhat differently, the lower the level of *APPROV*, the greater the *subjective benefits* ("secondary gain") from personal federal income tax evasion, whereas the higher the level of *APPROV* the lower the *subjective benefits* (secondary gain) of personal federal income tax evasion. Based on this symmetrical argument, it is hypothesized that the greater the public's approval rating of the President, the lower the *eb* and hence the lower the aggregate degree of personal income evasion, *ceteris paribus*. Accordingly, equation (5) is transformed into equation (6), as follows:

$$eb = j(ATR, TRA, PCTITEM, UN, APPROV), \quad (6)$$
$$j_{ATR} > 0,\ j_{TRA} < 0,\ j_{PCTITEM} > 0,\ j_{UN} > 0,\ j_{APPROV} < 0.$$

Finally, the second new variable integrated into this *time-series* framework is the interest rate yield on three year U.S. Treasury notes, *THREE*. This is a variable altogether overlooked in previous related empirical studies. It is argued here that the higher the level of *THREE*, the greater the expected benefits (*eb*) from engaging in income tax evasion since the dollars gained from that tax evasion can be invested in higher yielding securities. Alternatively stated, the higher the level of *THREE*, the greater the *opportunity costs of tax compliance*. Obviously, *THREE* is but one usable measure of the *opportunity costs of tax compliance*; for example, the yield on five year or ten year Treasury notes are reasonable alternative such measures. The adoption of THREE was based on the notion that it provides a greater yield than T-bills generally do, whereas it exposes its owner to much less interest rate risk than longer term notes (or bonds). Accordingly, it is hypothesized that the higher the value of *THREE*, the higher the *eb* associated with income tax evasion and hence the higher the aggregate degree of federal personal income evasion, *ceteris paribus.* Consequently, equation (6) is replaced by equation (7):

$$eb = j(ATR, TRA, PCTITEM, UN, APPROV, THREE), \quad (7)$$
$$j_{ATR} > 0,\ j_{TRA} < 0,\ j_{PCTITEM} > 0,\ j_{UN} > 0,\ j_{APPROV} < 0,\ j_{THREE} > 0.$$

The expected gross costs of not reporting income to the IRS are hypothesized to be an increasing function of the expected risks/costs thereof (Alm, Jackson, and McKee, 1992; Pestieau, Possen, and Slutsky, 1994; Erard and Feinstein, 1994; Caballe and Panades, 1997; Cebula and Coombs, 2009). In this study, to the representative economic agent, the expected risks/costs (*ec*) from not reporting or from underreporting taxable income to the IRS are enhanced by an increase in *AUDIT*, the percentage of filed federal personal income tax returns that is formally audited by IRS examiners, *ceteris paribus*. Indeed, the experience of an IRS tax audit could imply non-pecuniary ("psychic") costs as well as pecuniary costs (including outlays for legal or other representation, along with the value of one's own time) above and beyond any potential added taxes, penalties, and interest assessed by the IRS. In addition, to reflect further the risks associated with tax evasion, the variable *PEN* is included in the Model. *PEN* reflects the average total of interest and other penalties assessed by the IRS per audited tax return. Thus, we have:

$$ec = j(AUDIT, PEN),\ j_{AUDIT} > 0,\ j_{PEN} > 0. \tag{8}$$

Substituting from equations (7) and (8) into equation (1) yields:

$$\begin{gathered} pnr/(1\text{-}pnr) = b(ATR, TRA, PCTITEM, UN, APPROV, THREE, AUDIT, PEN), \\ b_{ATR} > 0,\ b_{TRA} < 0,\ b_{PCTITEM} > 0,\ b_{UN} > 0,\ b_{APPROV} < 0,\ b_{THREE} > 0,\ b_{AUDIT} < 0,\ b_{PEN} < 0. \end{gathered} \tag{9}$$

Let *AGI* represent the *actual total value of the aggregate federal adjusted gross income* in the economy, i.e., *AGI=UAGI+RAGI*, where *UAGI* is the dollar size of the *unreported aggregate federal adjusted gross income* in the economy, and *RAGI* is the dollar size of the *reported aggregate federal adjusted gross income* in the economy. It reasonably follows overall that:

$$UAGI = (pnr)*AGI \tag{10}$$

and

$$RAGI = (1\text{-}pnr)*AGI. \tag{11}$$

It then follows that:

$$UAGI/RAGI = (pnr)*AGI/(1\text{-}pnr)*AGI = (pnr)/(1\text{-}pnr). \tag{12}$$

From (9) and (12), substitution for *pnr/(1-pnr)* yields the following Model of aggregate personal income tax evasion:

$$\begin{gathered} UAGI/RAGI = b(ATR, TRA, PCTITEM, UN, APPROV, THREE, AUDIT, PEN), \\ b_{ATR} > 0,\ b_{TRA} < 0,\ b_{PCTITEM} > 0,\ b_{UN} > 0,\ b_{APPROV} < 0,\ b_{THREE} > 0,\ b_{AUDIT} < 0,\ b_{PEN} < 0. \end{gathered} \tag{13}$$

Empirical Analysis

Based on the framework provided in (13) above, the following reduced-form equation is to be estimated:

$$\begin{gathered} (UAGI/RAGI)_t = a_0 + a_1\, ATR_{t-1} + a_2\, TRA_t + a_3\, PCTITEM_{t-1} \\ + a_4\, UN_{t-1} + a_5\, APPROV_{t-1} + a_6\, THREE_{t-1} + a_7\, AUDIT_{t-1} + a_8\, PEN_{t-1} + u \end{gathered} \tag{14}$$

where:

$(UAGI/RAGI)_t$ = the ratio of the aggregate *unreported* federal adjusted gross income in year t to the aggregate *reported* federal adjusted gross income in year t, expressed as a percent;

a_0 = constant term;

ATR_{t-1} = the average effective federal personal income tax rate in year t-1, expressed as a percent;

TRA_t = a binary (dummy) variable for the years 1986 through 1987, when the Tax Reform Act of 1986 was initially implemented and became effective: $TRA_t = 1$ for the years 1986 and 1987, and $TRA_t = 0$ otherwise;

$PCTITEM_{t-1}$ = the percentage of filed federal personal income tax returns with itemized deductions listed on Schedule A of Form 1040 in year t-1;

UN_{t-1} = the percentage unemployment rate of the civilian labor force in year t-1;

$APPROV_{t-1}$ = the public's average job approval rating of the President in year t-1: values for $APPROV_{t-1}$ lie between 0 and 100;

$THREE_{t-1}$ = the average percentage interest rate yield on three year U.S. Treasury notes in year t-1;

$AUDIT_{t-1}$ = the percentage of filed federal personal income tax returns in year t-1 that was subjected to a formal IRS audit involving IRS examiners;

PEN_{t-1} = the average total of interest and other penalties assessed by the IRS per audited tax return in year t-1; and

μ = stochastic error term.

The study period runs from 1976 through 2005. The choice of the year 1976 reflects the limited availability of the itemized deductions data; the choice of the year 2005 reflects the most recent availability of the official *UAGI/RAGI* data. Naturally, this restriction implies that the number of observations is only 30 and the degrees of freedom in the various estimates provided here is in the range of 20. As a result, the criteria for statistical significance are commensurately higher than would be the case were more observations involved. The data are all annual. Following previous time-series studies of tax evasion using official data (Tanzi, 1982, 1983; Clotfelter, 1983; Carson, 1984; Long and Gwartney, 1987; Pyle, 1989; Feinstein, 1991; Erard and Feinstein, 1994; Feige, 1994; Cebula, 2001, 2004, 2008; Ali, Cecil, and Knoblett, 2001; Connelly, 2004; Christie and Holzner, 2006; Cebula and Coombs, 2009), the right hand-side variables (aside from the binary *TRA* variable) are lagged one year. This lagging not only is intended to minimize the possibility of simultaneity bias, but also to avoid specification bias that would result since the deadline for filing federal income tax returns is April 15th of each year and non-lagging would technically portray un-lagged variables as influencing past events. The *UAGI/RAGI* data were obtained from the IRS (2010, columns 2 and 3).The data for the variable *ATR* were obtained from the IRS (2009B). The *PCTITEM* data were obtained from the IRS (2009A). The *AUDIT* and *PEN* data were obtained from the Government Accounting Office (1996, Table I.1) and the U.S. Census Bureau (1994, Table 519; 1998, Table 550; 1999, Table 556; 2001, Table 546; 2009, Table 469). The *TRA* variable is a dummy variable. The data for the variables *THREE* and *UN* were obtained from the Council of Economic Advisors (2010, Tables B-73, B-35). The data for the variable *APPROV* were obtained from the Gallup Poll (2009). The (*P-P*) Phillips-Perron and *ADF* (Augmented Dickey-Fuller) unit root tests indicate that all of the variables in the Model are stationary in levels for the study period. The mean value for variable *UAGI/RAGI* for the study period was 13.42, with a standard deviation of 1.66.

The OLS (ordinary least squares) estimation of equation (14), adopting the Newey-West heteroskedasticity correction, is provided in Model 1 of Table 1. In Model 1, the coefficients on all eight of the explanatory variables exhibit the hypothesized signs, with six being statistically significant at the one percent level, one being statistically significant at the 2.5 percent level, and one being statistically significant at the five percent level.. The coefficient of determination is 0.89, so that the Model explains nearly nine-tenths of the variation in the dependent (tax evasion) variable. The *F*-statistic is significant at the one percent level, attesting to the overall strength of the estimate. Finally, with a *DW* = 1.90, there is no concern regarding autocorrelation.

Table 1. Empirical Estimates

	Model 1		***Model 2***		***Model 3***		***Model 4***	
Variable	***Coefficient***	***T stat***	***Coefficient***	***T stat***	***Coefficient***	***T stat***	***Coefficient***	***T stat***
Constant	15.2		15.3		8.69		25.9	
ATR	0.86	2.83**	0.85	2.61**	0.72	2.23*	0.86	2.58**
TRA	-4.16	-6.57***	-4.057	-4.06***	-4.62	-9.87***	-2.66	-4.31***
PCTITEM	0.21	4.41***	0.203	2.79**	0.306	8.37***	-------	-------
UN	1.066	6.13***	1.07	5.66***	0.94	5.96***	1.23	6.69***
APPROV	-0.173	-5.25***	-0.176	-4.00***	-0.109	-4.20***	-0.259	-7.45***
THREE	0.357	2.29*	0.35	2.22*	0.335	2.19*	0.316	1.63
AUDIT	-1.36	-4.07***	-1.345	-4.30***	-------	-------	-2.626	-5.06***
PENALTY	-0.49	-7.05***	-0.47	-3.75***	-0.60	-7.25***	-0.36	-3.03***
TREND	0.008	1.79#	-------	-------	-------	-------	-------	-------
R^2	0.89		0.89		0.86		0.83	
adjR^2	0.84		0.83		0.80		0.76	
F	17.57***		14.77***		16.25***		12.9***	
DW	1.90		1.88		1.78		1.79	
Rho	0.05		0.06		0.11		0.10	

***Statistically significant at 1% level; **statistically significant at 2.5% level; *statistically significant at 5% level; #statistically significant at the 10% level.

The estimated coefficient on the *ATR* variable is positive and statistically significant at beyond the two percent level. Thus, the higher the average federal personal income tax rate, the greater the degree of federal income tax evasion by households, presumably because a higher income tax rate increases the incentive to evade taxes. This finding is consistent in principle with the conventional wisdom and with several previous empirical studies, including Tanzi (1982), Clotfelder (1983), Feige (1994), and

Cebula and Coombs (2009). The estimated coefficient on the Tax Reform Act of 1986 dummy variable (*TRA)* is negative, as hypothesized (Musgrave, 1987), and statistically significant at the one percent level, providing evidence that taxpayers may have regarded the Tax Reform Act of 1986 as a genuine, honest effort to reform the inequities of and diminish the complexities (compliance costs) of the existing Internal Revenue Code. Alternatively, as implied by Slemrod (1992, p. 45), the observed drop in personal federal income tax evasion for this brief period (1986-1987) may simply have reflected the time frame required by taxpayers to learn about and adjust to this allegedly "sweepingly reformed" (Musgrave, 1987, p. 59) new version of the Internal Revenue Code. The estimated coefficient on the *PCTITEM* variable is positive and statistically significant at the one percent level, implying that the greater the percentage of federal income tax returns in which deductions are itemized on Schedule A of Form 1040, the greater the degree of federal personal income tax evasion. This finding is consistent with the panel data analysis of individual tax returns by Alm and Yunus (2009). The estimated coefficient on the unemployment rate variable (*UN*) is positive, as hypothesized, and statistically significant at the one percent level. This finding is consistent with the hypothesis that the higher the unemployment rate, the greater the degree to which households enter the underground economy (Alm and Yunus, 2009; Gahramanov, 2009; Cebula and Coombs, 2009). Next, there is the issue involving the Presidential job approval rating: "Does a lower (higher) job approval rating of the President by the U.S. public act to increase (decrease) the degree of aggregate federal personal income tax evasion?" As shown in Model 1 of Table 1, the estimated coefficient on variable *APPROV* is negative (as hypothesized) and statistically significant at the one percent level. Thus, this finding provides empirical support for this hypothesis (Cebula, 2008). As for the variable *THREE*, its coefficient is positive and statistically significant at the four percent level, implying (arguably) that the higher the interest rate yield on three year Treasury notes, the higher the opportunity costs of tax compliance.

The estimated coefficient on the variable *AUDIT* is negative (as hypothesized) and statistically significant at the one percent level. This finding would suggest that taxpayers are *discouraged* from tax evasion behavior by greater prospects of detection (as represented by variable *AUDIT* (Clotfelder, 1983; Feige, 1994; Cebula, 2008). Finally, the estimated coefficient on the variable PEN is negative and statistically significant at the one percent level, implying that the greater the IRS penalty plus interest assessments on detected unreported income, the greater the disincentive to engage in income tax evasion.

As tests of the robustness of the basic Model, three alternative versions of the basic Model have been estimated.. They are summarized in Models 2, 3, and 4 of Table 1. In Model 2, where a linear trend variable was added to the Model, *TREND*, the results very closely resemble those in Model 1. In Model 3 of Table 1, the estimate excludes the audit rate variable; in this case, the overall results for the remaining variables closely resemble their counterparts in Model 1 and 2. Finally, in Model 4, the variable *PCTITEM* has been deleted from the basic Model; although the estimated coefficient on variable *THREE* becomes statistically insignificant at the ten percent level, the results for the remaining variables largely resemble those in Model 1. Thus, the Model exhibits a reasonably high degree of consistency, i.e., robustness.

Conclusion

This study has used newly available data from the IRS (2010) on income tax evasion to identify key determinants of aggregate federal personal income tax evasion for the period 1976-2005. To date, only one related study has appeared that investigates beyond the year 1997, and that more narrowly focused study (Cebula and Coombs, 2009) runs only through the year 2001 and uses a non-IRS dataset (Ledbetter, 2004).

The empirical estimates provided in the present study indicate that the aggregate degree of federal personal income tax evasion, $(UAGI/RAGI)_t$, is *directly impacted by* the average effective federal personal income tax rate (*ATR*), the unemployment rate (*UN*), the interest rate yield on three year U.S. Treasury notes (*THREE*), and the percentage of filed federal personal income tax returns listing itemized deductions (*PCTITEM*). Aggregate federal personal income tax evasion also is negatively impacted by the variables *TRA* (reflecting the various provisions of the Tax Reform Act of 1986), *APPROV*, the public's job approval rating of the President *per se*, the percentage of filed tax returns formally audited by IRS examiners (*AUDIT*), and the IRS penalty assessment on detected unreported taxable income (*PEN*). The uniqueness of this study derives in part from the adoption of variables *PCTITEM* and *THREE*, which have not previously been analyzed in aggregate *time-series* studies of personal income tax evasion, whereas *none* of these factors has been investigated to date for the years 2002-2005, the most recent years for which the IRS has developed its newest estimates of federal household (personal) income tax evasion.

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A Panel Model of Branch Banking in the United States: 1994-2010

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Abstract

The last 25 years saw large-scale consolidation among banks and a rapid movement by banking institutions into other financial services. Consolidation points to downward pressure on bank branches, while new products increase the value of a branch system. This study assesses how these conflicting forces affected branch banking over the 1994-2009 period, the span since the passage of the Interstate Banking and Branch Efficiency Act. It uses a panel model where the 50 states serve as the panels. Economic, demographic, and market structure variables are all significant. Additionally, estimation results are consistent with observed trends in branch banking and enable one to separate the two competing forces affecting branch banking. Fixed effects, representing unique but unobserved variables in each state, are strong contributors to the model's explanatory power.

Introduction

Two major legislative initiatives in the 1990s set the stage for change in the years ahead. The Interstate Banking and Branch Efficiency Act (IBBEA) of 1994 [Federal Reserve Bank of Cleveland, 1966] enabled multi-bank holding companies to consolidate bank charters and made the formation of new multi-state banks far easier. Branch rationalization emerged as the bigger banks sought to improve earnings. The Financial Modernization Act (FMA) of 1999 [Federal Reserve Bank of Minneapolis, 1999, Furlong, 2000] streamlined the merger of banks and nonbank financial firms. This expanded the array of financial services that might be offered in bank offices and increased the importance of branches as conduits for the nonbank products.

Taking these factors into consideration, two opposing forces would be expected to surface. On the one hand, strong cost-control efforts, stemming from the bank consolidation process should lead to branch closings, downsizing of branches, and the outright sale of branches. Effects of these actions should produce a slowdown in the growth in the number of branches with an associated rise in the local-market population-to-branch ratio. Additionally, as branch growth slows, deposits-per-branch should rise for any given level of economic activity.

On the other hand, expanded business lines might encourage banks to maintain or even widen branch systems. Also, customers still need branches for traditional services, and consumer preference for branches remains strong [Costanzo, 2000; Fung, 2001; Habal, 2002; Jackson, 2002; Mandaro, 2002; Monahan, 2000]. For many, convenience is also a factor [Avery, Bostic, Calem, and Canner, 1997], and in branch efficiency studies, convenience is often cited as the justification for the seemingly large number of branches [Berger, Leusner, and Mingo, 1997; Cyree, Wansley, and Black, 2000].

Moreover, branches are becoming an important part of a multi-channel delivery system for a number of products. Banks are developing the branch delivery channel with large-scale technology spending [Rountree, 2002; Fung, 2002; Yulinsky, 2002]. There is also renewed recognition that branches are a key to building customer relations and to developing brand identity [Fung, 2001; Bills, 2002]. Bank of America may be the prototype for the shift that is underway. It had roughly 4,500 branches in 2000 [Mollenkamp and Beckett, 2000]. By 2002, it had been cut to roughly 4,000, but then planned on adding 200 branches a year beginning in 2004 [Bills, 2002]. Thanks to acquisitions, it had more than 6,300 branches by June 2006.

Several new strategies are underway, which are directed at generating more traffic and more business through existing branches. Some banks borrowed an idea from airport terminals and located on-line kiosks in branch offices on which a variety of activities can be undertaken [Wolfe, 2004]. Others are trying to tap the small business market more aggressively with in-branch small-business experts [Schmelkin, 2005]. Others are expanding into retail asset management and financial planning at the branch level [Morris, 2002].

With recognition of the rising value of a branch franchise, more customer-friendly branches began to appear in an effort to lure customers back into branch offices [Kuehner-Hebert, 2004]. Also, some have noted that banks should treat their branch system as a portfolio of branches, tailoring them within and between markets [Kaytes and Meleis, 2005].

As will be seen later in Table 1, the nationwide population-to-branch ratio declined slightly between 1994 and 2009, reflecting a slightly faster growth of branches than population. Separately, the ratio rose in 29 states, while 22 states and the District of Columbia witnessed a slower growth in population than in branches over the same period (Table 2, shown later). On the surface, this finding suggests that positive forces noted above slightly dominated negative forces on branch growth. This study analyzes these developments in light of the competing forces noted above. To do this, it identifies the determinants of the local market structure of banking and uses that information to assess effects of the banking laws and general competition forces on local banking markets. Local market structure is defined by two variables: (1) the number of banking offices in the market and (2) the intensity with which each bank is utilized as measured by the average branch size in local markets.

It is likely that the effect of economic and demographic factors on branches and deposit footing will vary across local markets. To assess this possibility, while still focusing on the nation as a whole, this study defines the local markets as the 50 states and the District of Columbia. Given the presumption that differences exist among markets, a longitudinal study is planned in which the local markets serve as panels. This approach allows the development of 51 panels, encompassing the 50 states plus the District of Columbia.

Admittedly, states may represent an over-sized local market. However, there is some evidence that states are becoming the relevant market, given the industry's deregulation. Morgan (2006) found that branch prices are related to statewide data rather than local data for the northeastern states. Deregulation allowed the market for bank services to widen, and hence price came to be based more on statewide than local bank conditions. In any event, the state level of aggregation does serve the purpose of this study. In terms of the study's structure, attention now turns to the existing research on branch banking, followed by the study's model, the estimation results, and concluding remarks.

Earlier Research

Research at the branch level typically focuses on branch efficiency. As such, those studies take the existing structure as given and offer no basis for the existing structure. Some, however, have broken from that typical approach. Chang, Chaudhuri and Jayaratne (1997) found evidence of a herding effect in the New York City market. Additionally, effects of herding on branch numbers were at odds with economic and demographic variables in the local markets.

Avery (1991) looked to five city-wide markets and their respective branch structure in 1977 and 1988-1989. That study found that consolidation did not have an adverse effect on branch structure in low income areas, when taking account of demographic factors. Also, between those years, branches per capita rose in those markets, despite the fear that deregulation would cut branches. Later, Avery, Bostic, Calem, and Canner (1999) found that consolidation led to a systematic elimination of branches but mergers led to a branch overlap within zip-codes. Once a wider market is considered, mergers did not necessarily lead to a fall in branches relative to population. Where effects were felt, they were more serious in low income areas.

The studies assessing effects of consolidation on branch structure look to market-wide data, though the definition of the relevant market may vary from study to study. The decision to open or close a branch is based upon the expected performance of that branch; yet the studies evaluating consolidation ignore the profitability-based decision making process. Nonetheless, banks have proprietary models to assess branch profitability, which have a heavy bearing on decisions to open or close branches. Unfortunately, little is available on the propriety model; however, Avkiran (1997) fills this void, developing a model using "major business drivers" which allows for the comparison of actual and predicted performance.

As with the consolidation studies, Avkiran takes the existing structure as given. It offers no explanation as to the economic rationale for the overall relationship between branches and economics, except to the extent that local economic conditions are incorporated into the business drivers. It, along with the consolidation studies, seems to presume that the overall structure is irrelevant. Thus, there remains the unanswered question as to what influences the overall number of branches and the average size of branches in local markets. This study fills that void.

The Model

The starting point in the development of the model used to study the branch structure is a simple demand for money function such as those used by Dwyer and Hafer (1999), Lowan, Peristiani, and Robertson (1999), Mehra (1997), and Moghaddam (1997) and denoted in log form as

$$\log\left(\frac{M}{P}\right) = \beta_0 + \beta_1 \log(y) \tag{1}$$

However, this study replaces real money balance with nominal deposits. The model then links the nominal deposits (*dep*) for any period (*t*) to the nominal level of personal income (*pi*) and other factors (*of*). Each year (*t*) is composed of *j* observations, which represent the nation's 50 states and the District of Columbia. The states serve as the *j* panels in this study for *j* = 1 … 51. This may be represented in general form as

$$dep_{j,t} = F\left(pi_{j,t}, of_{j,t}\right) \tag{2}$$

where j = the j^{th} state, and $j = 1, ..., 51$ and
t = the t^{th} year.

The study next assumes that the supply of deposit services (*dep*) in local markets can be provided in one of two ways (1) the number of branches (*br*) and (2) the size of the branches expressed as deposits-per-branch (*dep/br*). In equilibrium, the quantity of deposit services demanded equals the quantity supplied, and that equilibrium condition may be expressed as

$$dep_{j,t} = br_{j,t} \times \frac{dep_{j,t}}{br_{j,t}} = F\left(H, G\right) \tag{3}$$

This approach allows for the separate analysis of the number of branches and the deposits per branch per state as functions of income and other factors. This may be denoted by

$$br_{j,t} = G\left(pi_{j,t}, of_{j,t}\right) \tag{4}$$

and

$$\frac{dep_{j,t}}{br_{j,t}} = H\left(pi_{j,t}, of_{j,t}\right) \tag{5}$$

Personal income is the product of (1) demographic forces represented by population (*pop*) and (2) economic forces represented by per capita income (*pi/pop*). To capture these separately, personal income (*pi*) is disaggregated into *pop* and *pi/pop*.

Other factors (*of* from Equation 1) may also influence a state's branch structure. A banking institution's willingness to establish branches in a state may be influenced by its income distribution (*dist*) in addition to per-capita income. Population density (*density*) may also affect branches. States with dense population may be more desirable than states with sparse population. Growth may be more predictable in densely populated states than in rural states, and denser population may make it easier for banks to gain market share than in thinly populated states.

It is unlikely that changes in economic or demographic conditions lead to a contemporaneous change in branch structure. As a result, it was decided to lag the economic and demographic data one year compared with the branch data. Admittedly, a longer distributed lag may be operative, but the one-year lag is considered a reasonable approximation of the process.

Expanding the general form of the supply of branch function is expressed in its level form as

$$br_{j,t} = b_{0,j}\left(pop_{j,t-1}\right)^{b_{1,j}} \times \left(\frac{pi_{j,t-1}}{pop_{j,t-1}}\right)^{b_{2,j}} \times \left(dist_{j,t-1}\right)^{b_{3,j}} \times \left(density_{j,t-1}\right)^{b_{4,j}} \times \left(conc_{j,t}\right)^{b_{5,j}} \tag{4-1}$$

As with the simple money demand functions, this function will be estimated in logs, and the function may then be expressed as:

$$log(br_{j,t}) = b_{0,j} + b_{1,j} log(pop_{j,t-1}) + b_{2,j} log\left(\frac{pi_{j,t-1}}{pop_{j,t-1}}\right) + b_{3,j} log(dist_{j,t-1}) + b_{4,j} log(density_{j,t-1}) + b_{5,j} log(conc_{j,t}) \tag{5}$$

A similar function is used to explain the deposit-per-branch ratio.

$$log(dep_{j,t}/br_{j,t}) = b_{0,j} + b_{1,j} log(pop_{j,t-1}) + b_{2,j} log\left(\frac{pi_{j,t-1}}{pop_{j,t-1}}\right) + b_{3,j} log(dist_{j,t-1}) + b_{4,j} log(density_{j,t-1}) + b_{5,j} log(conc_{j,t}) \tag{6}$$

The states (j) are the panels as well as the cross-sectional identifiers. Similarly for $b_{1,j}$ through $b_{5,j}$, there will be a separate coefficient calculated for each state in the cross-section version for a total of 51 coefficients on each cross-sectional variable. If the economic and demographic variables are found to have only common effects, the 51 coefficients collapse into one for that variable. Finally, both fixed effects and random effects methods are considered in the estimation of $b_{0,j}$ coefficients.

The Data

This study uses data on the number of deposit-taking offices and total deposits by state of banks and savings institutions insured by the Federal Deposit Insurance Corporation (henceforth called the FDIC). Banking offices include the main and branch offices of each institution in the local market. Data are from the FDIC's online databases and are annual observations as of June 30 of each year. The data used in this study are for June 1994 through June 2009. This span represents the extent of the data available on the FDIC web site. Population and income data are from the Bureau of Economic Analysis (BEA) online databases. Population density is measured as population per square mile (*pop/sqmi*) for each state. Per capita dividends and per capita transfer was used as proxies for income distribution (*dist*) these data are also available through the BEA on-line databases. Finally, concentration is measured by the statewide Herfindahl-Hirschman Index (*HH*I) which is available through the FDIC website.

Empirical Results

Broad Trends

Table 1 reports the trends in the overall branch and deposit structure as well as the income and demographic variables over the 1994-2009 span. In terms of branch developments, there was an average of one branch for every 3,217 persons nationwide in 1994 (line 8) compared to 3,106 persons in 2009. This stemmed from a slightly faster growth of branches of 1.36 percent (line 2) compared with population growth of 1.12 percent (line 8), suggesting that business opportunities of branch growth slightly dominate cost-control initiatives. This led to a branch-to-population elasticity for branches of 1.21 percent in which each 1 percent population growth is associated with a 1.21 percent growth in branches. Numerically, each new branch was associated with a population increase of 2,594 persons, on average, between 1994 and 2009 across the 50 states and the District of Columbia.

Effects were mixed across states. As seen in Table 2, 29 of 51 states posted a branch-to-population elasticity measure greater than one, while 22 registered an elasticity measure less than 1. Differences in branch to population elasticities are likely attributable to the various economic measures in the model developed below. While differences are evident among states, the dominance of elasticity measures greater than one is consistent with the overall nationwide branch-to-population elasticity measure of 1.21.

Table 1: Nationwide Trends, 1994-2009

		1994	2009	Percent Change*
	Branches and Deposits			
1	Total Deposits (Billions of $)	$3.132	$7.493	5.99
2	Branches	80788	98,943	1.36
3	Average Branch Size (Millions of $)	$38.769	$74.929	4.49
	Economic and Demographic Data			
4	Disposable Personal Income (Billions of $)	$5.144	$10.915	5.14
5	Per-Capital Disposable Income (Thousands of $)	$19.791	$35.553	3.98
6	Population (Millions of Persons)	259.919	307.007	1.12
	Deposit Richness			
7	Deposit-to-Disposable Personal Income Ratio	0.61	0.69	0.80
8	**Population and Branches**			
	Population-to-Branch Ratio	3217	3103	-0.24
	Income Distribution			
9	Per Capita Dividends, Interest, and Rents	$0.996	$6.708	13.56
10	Per Capita Transfer Payments	$0.790	$6.868	15.51

* At a compound annual rate

One explanation of the strong nationwide elastic response of branches relative to population might lie in the proliferation of branches in supermarkets [Agosta, 2000; Bach, 2000a; Bach, 2001, Boraks, 2002; Silvestri, 2000; Winokur, 1999a, 1999b]. This trend continues even at the present [Davis, 2005b]. Some see such branches becoming an increasingly important distribution channel in an ever more competitive banking environment with the convenience store as the next site for a limited-service presence [Tescher, 2005]. Nonetheless, supermarket branches are not always successful [Quinn, 2000; Winokur, 2000], and these branches tend to top out sooner than conventional branches and lack the ability to provide a broad range of products. This led one large bank to shutter in-store branches in favor of physical branches in one of its markets [Davis, 2005a]. Supermarket branches also have a downside risk. If a chain closes a large number of stores in which a bank has branches, an important delivery vehicle is lost to that bank—as in the case when Albertson closed its stores [Reosti, 2002].

In terms of deposits, cross-industry competition may have had an effect. If nonbank financial firms siphon deposit-type business from banking institutions, deposits per branch should grow more slowly than personal income during the sample period. Additionally, since forces of expansion seem marginally stronger than cost rationalization, the marginally stronger growth in the number of branches could retard the growth of average branch size. Interestingly, this is borne out in the aggregated data. Deposits-per-branch climbed at a 4.49 percent rate, while disposable personal income rose at a 5.14 percent rate.

Screening Results

Branches per state (Equation 5) and the deposits per branch (Equation 6), was estimated over the 1994-2009 period. Screening results showed that all the relevant variables had a common effect on both dependent variables. Coefficients were generally statistically significant, and in most cases, at least at the one percent level. However, deposit richness was not significant and reported results do not include it. Similarly, population density was dropped, largely due to its marginal statistical significance in both the branch function and the deposits per branch function.

Table 2: Branch-to-Population Elasticities

	Statewide Elasticity	Elasticity Greater than 1		Statewide Elasticity	Elasticity Greater than 1
AK	0.349		MT	3.891	YES
AL	1.750	YES	NC	0.274	
AR	3.147	YES	ND	64.070	YES
AZ	1.162	YES	NE	2.862	YES
CA	0.579		NH	0.609	
CO	2.693	YES	NJ	1.362	YES
CT	0.718		NM	0.826	
DC	-2.069		NV	0.961	
DE	0.336		NY	1.999	YES
FL	1.170	YES	OH	0.723	
GA	0.936		OK	3.302	YES
HI	-3.014		OR	1.170	YES
IA	3.222	YES	PA	1.959	YES
ID	1.269	YES	RI	-1.033	
IL	5.772	YES	SC	0.829	
IN	0.932		SD	2.475	YES
KS	3.438	YES	TN	1.473	YES
KY	2.120	YES	TX	2.426	YES
LA	5.054	YES	UT	0.798	
MA	1.722	YES	VA	0.618	
MD	0.254		VT	0.354	
ME	0.968		WA	0.460	
MI	1.217	YES	WI	2.256	YES
MN	2.103	YES	WV	-357.514	
MO	2.603	YES	WY	5.464	YES
MS	1.231	YES			

The Panel Results

Results for the branch function is reported in Table 3 (Model 1 represents the fixed effects versions and Model 2 the random effects version. Table 5 reports results for deposits per branch with Model 3 representing fixed effects and Model 4 random effects. Before proceeding any further, this brings the study to the issue of fixed versus random effects. The study's panel model examines the unobserved factor unique to each state that affects the number of branches (Equation 5) and deposits per branch in each state. Either fixed effects or random effects may be used to represent effects unique to each state but not captured elsewhere. The fixed effects coefficients capture collective effects of unobserved and unidentified cross-sectional variables that are time invariant. Random effects capture effects that may be time variant.

Here, the Hausman (1978) test was used to determine whether fixed or random effects were appropriate. The null hypothesis (no statistical difference between the fixed and random effects parameter estimates) was rejected with a p-value of close to zero (<0.001). From this, it appears that the estimates from the random effects model are not efficient compared with the fixed effects model, and the fixed effects model is preferred. Additionally, a test for the redundancy of the fixed effects terms was also conducted. For both the F-test and the Chi square test for redundancy, the null was rejected with a p-value of <0.001, implying that the fixed effects contain information not captured elsewhere in the model. In other words, there remains unidentified information unique to each institution that is time invariant. Table 4 reports the fixed effects coefficients for both the branch function (Equation 5) and the deposits per branch function (Equation 6).

Thus, while results for both fixed and random effects are reported in Tables 3 and 5, the discussion is confined to the fixed effects versions. Estimates of the fixed effects are reported in Table 4 for both the deposits and the deposits per branch function. The discussion now turns to the deposit function.

Branches per State

Estimation results for branches per state are reported in Model 1 (Table3). Results are generally as expected. Population and per-capita disposable personal income behave as expected, and both coefficients differ from zero with virtual certainty. Similarly, effects of income distribution are as expected. Per capita dividends have a positive sign, while transfer payments have a negative effect. However, the dividends coefficient differs from zero with virtual certainty, but the transfers coefficient lacks significance, except at a low level (15 percent) of confidence. The positive effect of per capita dividends suggested that "richness," as measured by per capita dividends, encourages branch formation. For transfer payments, the negative coefficient suggests that poorness discourages branch formation, but its relevancy is open to question given the coefficient's lack of statistical significance. Finally, as expected, concentration plays a role. High concentration discourages branch formation and vice versa.

Table 3: Estimation Results 1994-2009

Model 1: Log(Branches) and Fixed Effects			**Model 2: Log(Branches) and Random Effects**		
Variables	**Coef.**	**P-Values**	**Variables**	**Coef.**	**P-Values**
C	-4.990	0.000	C	-6.754	0.000
LOG(POP(-1))	0.763	0.000	LOG(POP(-1))	0.892	0.000
LOG(DPI(-1)/POP(-1))	0.201	0.006	LOG(DPI(-1)/POP(-1))	0.165	0.018
LOG(DIV(-1)/POP(-1))	0.090	0.008	LOG(DIV(-1)/POP(-1))	0.077	0.024
LOG(TRAN(-1)/POP(-1))	-0.093	0.154	LOG(TRAN(-1)/POP(-1))	-0.072	0.231
LOG(HHI(-1))	-0.017	0.005	LOG(HHI(-1))	-0.027	0.000
Cross-section fixed effects			Cross-section random effects		
51 coefficients displayed in Table 4			51 coefficients estimated but not displayed		
Summary Statistics			*Summary Statistics*		
Periods	16		Periods	16	
Cross-sections	51		Cross-sections	51	
Total Observations	806		Total Observations	806	
Adjusted R-squared	0.997			S.D.	Rho
F-statistic	4629.672		Cross-section random	0.163	0.898
Prob(F-statistic)	0.000		Idiosyncratic random	0.055	0.102
			Weighted Statistics		
			Adjusted R-squared	0.778	
			F-statistic	565.779	
			Prob(F-statistic)	0.000	
			Mean dependent variable	0.594	
			S.D. dependent variable	0.122	
			Unweighted Statistics		
			R-squared	0.936	
			Mean dependent variable	7.032	

Notes: Fixed effects estimated by panel least squares with White cross-section standard errors and covariance with degree of freedom correction. Random effects estimated by panel EGLS with White cross-section standard errors and variances with degree of freedom correction and Swamy and Arora estimator of component variances.

In reviewing the results, the coefficient (0.763) on population is less than one, implying that the branch-to-population ratio would fall given a slower growth in branches compared with population implied by the coefficient. However, population appears elsewhere in the function, and taking that into account, the elasticity of branches with respect to population is the derivative of the branch function in Model 1with respect to population which becomes

$$\frac{d\left(\log\left(br_t\right)\right)}{d\left(\log\left(pop_{t-1}\right)\right)} = 0.763 - 0.201 - 0.090 - (-0.093) = 0.565 \qquad (7)$$

In other words, each percentage point change in population leads to only a 0.5659 percentage point change in the number of branches, which is still less than unity. This finding is consistent with the view that consolidation and cost rationalization leads to a more parsimonious growth in branches compared with population. At the same time, the estimation results may seem inconsistent with the observed elasticity measure of 1.16 percent over the 1994–2009 period. This difference is likely attributable to the other explanatory variables in the model, and they also have a bearing on the overall percent change in branches. Thus, the growth in disposable income, dividends, and transfer payments has a net effect that pushes the observed population elasticity above one. Thus, the strength provided by positive economic growth offsets, on average, the restraining effect of consolidation and cost controls on branch growth.

Fixed Effects

As can be seen (Table 4), there is not a wide variability in the fixed effects coefficients among the panels (the states) in the branch function (Table 3, Model 1).

Table 4: Estimate of Fixed Effects

Model 1: Log(Branches)				**Model 3: Log(Deposits/Branches)**			
AK	-0.862	MT	-0.198	AK	3.506	MT	2.871
AL	0.078	NC	0.162	AL	-0.851	NC	-2.154
AR	0.338	ND	0.255	AR	0.368	ND	3.588
AZ	-0.417	NE	0.328	AZ	-1.169	NE	1.443
CA	-0.047	NH	-0.268	CA	-4.952	NH	2.337
CO	-0.139	NJ	0.244	CO	-0.516	NJ	-1.952
CT	-0.030	NM	-0.316	CT	0.045	NM	1.156
DC	-0.416	NV	-0.590	DC	4.330	NV	1.631
DE	-0.394	NY	0.119	DE	5.154	NY	-3.399
FL	0.212	OH	0.336	FL	-3.601	OH	-2.826
GA	0.074	OK	0.049	GA	-2.038	OK	-0.037
HI	-0.504	OR	-0.149	HI	2.478	OR	-0.421
IA	0.385	PA	0.418	IA	0.321	PA	-3.113
ID	-0.163	RI	-0.651	ID	1.627	RI	2.906
IL	0.251	SC	0.027	IL	-2.515	SC	-0.738
IN	0.256	SD	0.172	IN	-1.523	SD	3.600
KS	0.366	TN	0.219	KS	0.470	TN	-1.413
KY	0.287	TX	0.110	KY	-0.567	TX	-4.069
LA	0.137	UT	-0.371	LA	-0.981	UT	1.563
MA	0.066	VA	0.153	MA	-1.163	VA	-1.729
MD	0.009	VT	-0.122	MD	-1.149	VT	3.566
ME	-0.026	WA	-0.016	ME	1.831	WA	-1.528
MI	0.169	WI	0.273	MI	-2.700	WI	-1.076
MN	0.050	WV	-0.014	MN	-0.950	WV	1.209
MO	0.243	WY	-0.415	MO	-1.202	WY	4.531
MS	0.217			MS	-0.057		

All are fairly evenly tightly distributed about zero, but one mild outlier is evident— Alaska. Its negative sign could be attributed to the state's the low population and wide geographic area which leads to fewer branches than in more compact markets.

The more interesting finding is the negative sign for all states when the constant term (-4.99) from Table 3 (Model 1) and the fixed effects coefficients are added together. This suggest that the unobserved variables have an overall negative influence on branches per state, and the constant and fixed effects coefficients could (collectively) be the model's estimate of the average effect of branch rationalization and cost controls over the sample period. Thus, their effects are captured in the estimation results, notwithstanding the seemingly stronger growth in branches than population in the overall data.

Deposits per Branch per State

Results for the average branch size by state are reported in Table 5 for fixed effects (Model 3) and random effects (Model 3). The focus is again on the fixed effects variant. Results are generally as expected. In this version, population and per capita personal income are hypothesized to be key demographic and economic determinants of average branch size. As with the branch function, both coefficients differ from zero with virtual certainty.

Table 5: Estimation Results 1994-2009

Model 1: Log(Deposits/Branches) and Fixed Effects			**Model 4: Log(Deposits/Branches) and Random Effects**		
Variables	**Coef.**	**P-Values**	**Variables**	**Coef.**	**P-Values**
C	-34.593	0.000	C	-3.718	0.000
LOG(POP(-1))	2.375	0.000	LOG(POP(-1))	0.250	0.000
LOG(DPI(-1)/POP(-1))	0.214	0.076	LOG(DPI(-1)/POP(-1))	0.396	0.024
LOG(DIV(-1)/POP(-1))	-0.112	0.141	LOG(DIV(-1)/POP(-1))	0.078	0.350
LOG(TRAN(-1)/POP(-1))	0.265	0.054	LOG(TRAN(-1)/POP(-1))	0.373	0.032
LOG(HHI(-1))	0.278	0.000	LOG(HHI(-1))	0.283	0.000
Cross-section random effects:			Cross-section random effects		
51 coefficients displayed in Table 4			51 coefficients estimated but not displayed		
Summary Statistics			*Summary Statistics*		
Periods	16		Periods	16	
Cross-sections	51		Cross-sections	51	
Total Observations	806		Total Observations	806	
Adjusted R-squared	0.886			S.D.	Rho
F-statistic	282.821		Cross-section random	0.343	0.791
Prob(F-statistic)	0		Idiosyncratic random	0.177	0.209
			Weighted Statistics		
			Adjusted R-squared	0.588	
			F-statistic	228.326	
			Prob(F-statistic)	0.000	
			Mean dependent variable	0.491	
			S.D. dependent variable	0.303	
			Unweighted Statistics		
			R-squared	0.336	
			Mean dependent variable	3.826	

Notes: Fixed effects estimated by panel least squares with White cross-section standard errors and covariance with degree of freedom correction. Random effects estimated by panel EGLS with White cross-section standard errors and variances with degree of freedom correction and Swamy and Arora estimator of component variances.

The proxies for income distribution have an opposite effect on deposits per branch compared with the results from the branch equation. Per-capita dividends have a negative effect, while and per capita transfers have a positive effect. In contrast to Model 1, (branches), the dividends' coefficient differs from zero with marginal certainty (p=0.145), while transfer have a more reasonable level of significance (p=0.054). Model 3 results suggest that branches in wealthy states (i.e., states with high per capita dividends) will likely be smaller than branches in less wealthy states, other things equal. This is likely the result of wealthy individuals having a greater variety of alternatives to bank deposits investment and transaction vehicles than less wealthy individuals. The opposite effects of dividends in the two models should not be surprising. With the positive effect of per capita dividends on branches, effects would likely be opposite for average branch size. If richness leads to more branches, then it should have a diluting effect on the average size of the branches.

Turning to per capita transfer, a state with a higher share of low income residents, as measured by per-capita transfer payments, leads to higher average deposits than states with lower per-capita transfer payments. Lower-income individuals have few alternatives to bank deposits as investment and transaction vehicles, and hence states with high per-capita transfer payments are expected to have higher average deposits, other things equal. Finally, and not surprisingly, concentration adds to average deposits per branch. As with effects of income distribution, if concentration discourages branch formation, it should also enhance average deposits per branch.

Next, the growth of average deposits per branch can be expressed as the derivative of the function with respect to time. This may be denoted by

$$\frac{d\left(\log\left(\frac{dep_{j,t}}{br_{j,t}}\right)\right)}{d(t)} = 2.375\times\frac{d\left(\log\left(pop_{j,t-1}\right)\right)}{d(t)} + 0.214\times\frac{d\left(\log\left(\frac{pi_{j,t-1}}{pop_{j,t-1}}\right)\right)}{d(t)} - 0.112\times\frac{d\left(\log\left(\frac{div_{j,t-1}}{pop_{j,t-1}}\right)\right)}{d(t)} + 0.265\times\frac{d\left(\log\left(\frac{trans_{j,t-1}}{pop_{j,t-1}}\right)\right)}{d(t)} + 0.278\times\frac{d\left(\log\left(HH\right)\right)}{d(t)} \tag{8}$$

For positive growth in all the explanatory variables, average deposits per branch will exhibit positive growth. In times of recession, effects on the growth of average assets per branch depend upon the relative magnitudes of any variables that may decline. Since the demographic and economic variables increased over the 15-year sample period, they account for the rise in average deposits per branch. Thus, the model's results are consistent with observed behavior.

The fixed effects account for unique differences among the states. Again, the fixed effect coefficients are distributed about zero, but not as tightly as in Model 1. Several large positive values stand out—Delaware (due possibly to special purpose banks), and the District of Columbia, Wyoming, North Dakota, and Alaska. Large negative factors are observed in California and Florida. When combined with the constant term, they are negative. The implication is that unobserved variables have a restraining effect on branch deposit state by state, offsetting effects of the economic and demographic variables. Economically, they could reflect the competitive drain on bank deposits from nonbank institutions.

Epilogue

Several important findings stand out. A longitudinal model of branches and deposits per branch can be successfully developed at the state level utilizing states as the panels. In that model, population, per-capita income, income distribution, and concentration play key roles in determining the number of branches per state and average deposits per branch per state. All have common effects across states for both branches and average deposits per branch per state.

In terms of income distribution, a state's richness encourages branch formation, but restrains deposits per branch. A state's relative poorness plays has an opposite effect, though its lack of statistical significance suggests that richness may play a more prominent role than relative poorness in strategic business decisions. The opposite holds for deposits per branch, relative wealth reduces average deposits, while relative poorness enhances average balances. However, given the statistical insignificance of per-capita transfers, it appears that relative poorness plays a less systematic role than relative richness across states in terms of deposits.

In terms of the basic hypothesis laid out at the start of the paper, the model's estimate of the population elasticity for branches is less than one which is consistent with view that competitive pressure stemming from bank consolidation restrains the growth in branches. The faster observed branch growth relative to population is thus attributable to the power of economic effects over rationalization pressures. Also, evaluation of the time derivative of average branch size is positive, except for very large declines in the explanatory variables, which is consistent with its pattern over the sample period.

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A Note on Rating Implications of CDO for the Originating Bank's Market Value

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Abstract

Rating grades have a severe impact on securitizations as they determine regulatory capital requirement. In theory, risk is shifted from the originator to the bankruptcy-remote special purpose vehicle (SPV) that issues the securitization. Hence, the issuing bank's shareholder should not bother about rating announcements of securitizations. We will focus on securitizations such as collateralized debt obligations (CDO) and present that rating announcements only have a significant wealth effect on the originating bank around the announcement day. We will highlight that due to the severe mistrust to CDO in the financial crisis market participants no longer relied on rating grades and have already incorporated upcoming rating announcements for the originating bank.

Introduction

Asset Backed Securities (ABS) as the early form of securitization has seen a rapid growth since the 1990s. Later forms of securitizations such as Collateralized Debt Obligation (CDO) have extensively emerged and account for a large part of the issuing volume of securitizations. In 1996 the outstanding volume of ABS accounted for $456 billion and grew to $2.8 trillion in 2006 with CDO soaring to $550 billion in 2006 (Higgins et al., 2009).

Along with the increasing significance of securitizations, credit rating agencies' importance grew as their rating grades determine regulatory capital requirement. With structured credit products like securitizations drawing attention to investors, Mason and Rosner (2007) provide evidence that rating agencies evolved more towards underwriters. Due to their in-depth market knowledge, investors relied on the rating grades as a quality measurement. Nonetheless, a rating is not an exact measure of the default risk but rather eases comparisons across issuers by means of categorized risk measures. S&P, Moody's and Fitch - the three largest rating companies accepted by the supervisory authorities – provide ratings to a broad range of issuers and products. Each of the three agencies has its own rating grade category. Despite these separate risk categories, investors and the rating agencies themselves have a common understanding of the various risk categories and methodologies. There are some differences in the methodologies across the rating agencies. Moody's considers the default risk and the expected recovery rate whereas S&P and Fitch take the default risk into account. Rating agencies provide long-term and short-term ratings. We consider long-term CDO deals and therefore focus on long-term ratings.

Supervisory institutions restrict the investment opportunities of institutional investors to a certain rating threshold. Backed by favorable legislation according to regulatory capital requirement, high yielding CDO deals with investment grade status became attractive investment opportunities. However, Skreta and Veldkamp (2009) show that rating agencies underestimated risk and provided better than justified ratings since providing rating grades to CDO deals became a profit contributor. Rating grade inflation came in line with the growing significance of securitization in general and CDO in particular. First market distortions were noticed in 2006 with high default rates of subprime mortgages in the USA. Rating agencies announced upcoming warnings and downgrades. Most of the downgrades took place in the home equity ABS market but resulted in mistrust to securitizations in general (Ashcraft and Schuermann, 2010). As a result of more restrictive rating grade approval, many rating grades of CDOs have seen a dramatic downgrade and became less attractive to investors.

In theory, rating announcements of securitizations should not affect the originator. In order to issue CDO, the originator founds a special purpose vehicle (SPV) that is bankruptcy-remote of the originating institution. We consider financial institutions as originators in this paper. CDO are constructed so that defaults of the underlying portfolio result in losses for the originator by absorbing the retained first-loss-piece or equity tranche. During the financial crisis it was perceived that banks were not prepared to absorb losses because they did not possess that much equity on-balance. If losses are not absorbed by the first loss piece, losses are directed to the mezzanine tranches and finally to the senior tranches. These risk mitigation techniques represent an implicit recourse and, hence, may have an impact on the originating bank when rating downgrades occur (Higgins and Mason, 2004; Duffee, 2009; Ivashina, 2009).

With the theoretical background in mind, we provide evidence that rating announcements of CDO - mostly downgrades or negative outlooks in recent years - have indeed a negative impact on the issuing bank's share price around the announcement day but are generally incorporated as CDO suffered severe market mistrust. We conduct with section 2 summarizing the relevant literature, section 3 proposes the applied methodology and data, section 4 presents the results and section 5 finally concludes the research.

Literature Review

Implications of rating announcements have been researched in various ways either focusing on the equity or bond market or even both. Generally, if rating announcements convey new information, prices should react to the new information. Higgins et al. (2009) document that rating downgrades of ABS as the superior form of securitizations induce significant negative share price reactions and future securitization issuing cycles are affected by rating downgrades. Furthermore, rating agencies take the financial situation of the originating bank into account which suggests that investors are well aware about the quality of the bank they invest in. Our paper concerns about CDO as a sub-form of ABS. In a CDO the reference pool can consist of already securitized products such as ABS. Hence, it is less transparent to reveal the ultimate parent institution when examining CDO (Ammer and Clinton, 2004). We focus on CDO where rating implications have not yet been widely researched.

As the reference portfolio of a CDO deal is first pooled to a SPV and then issued as bonds to investors, we will conduct the literature review with rating implications on bonds. Pinches and Singleton (1978) found that bond rating changes were anticipated so that share price response did not reveal abnormal returns after the announcement. Contrary results were found by Katz (1974) resulting in delayed abnormal negative performance after downgrades. Grier and Katz (1976) conducted a survey on bonds from utilities and industrials. Anticipation was only perceived for industrials and price changes were stronger after downgrades. Early anticipation with no abnormal returns were investigated by Weinstein (1977). Other research analyzed weekly abnormal bond returns and found significantly negative returns in the week of downgrades but no abnormal returns for upgrades (Wansley et al., 1992). In the findings of Hite and Warga (1997) downgrades were preceded by negative abnormal returns. Steiner and Heinke (2001) conducted a study including not only downgrades but also negative watch listings. They found significant negative abnormal returns well before downgrades and negative watch listings and evidence for overreaction right after the event date.

Besides bond price implications we continue the literature review with the effect on equity prices. Academic researchers have gained proof that rating downgrades have more impact on share prices than rating upgrades (Griffin and Sanvicente, 1982; Holthausen and Leftwich, 1986; Glascock et al., 1987). Not necessarily do all rating downgrades have a negative impact. Investors distinguish between anticipated news and news that are based on deteriorating financial figures (Goh and Ederington, 1993). Rating downgrades that are driven by slurping financial figures show a significant negative impact whereas positive announcements result in rising equity prices. Despite this generalization, not all negative rating announcements are succeeded by negative share prices. For instance, the increase in the leverage ratio shifts wealth from debt to equity holders by generating higher expected returns to equity holders. In case of CDOs, the majority of the rating announcements is backed by deteriorating assets in the reference pool. Therefore we would expect falling share prices from negative rating announcements. We will challenge if this assumption is valid since the founding of a bankruptcy-remote SPV transfers risk from the originating bank to the SPV.

Cross-sectional differences based on firm and issue characteristics were observed in different publications. Researchers gained knowledge of share and bond price decline from rating downgrade announcements (Hand et al., 1992; Dichev and Piotroski, 2001). With credit default swaps gaining importance in recent times, (Hull et al., 2004; Norden and Weber, 2004) conclude that the reaction of credit default swap prices is most pronounced for rating reviews for downgrade. Despite finding significant results for rating downgrades or reviews for downgrades, recent research concludes that the market anticipates rating announcements. Hence, price effects take effect long before the actual announcement date. (Covitz and Harrison, 2003) surveyed that ca. 75% of the change in bond spreads occurs in the six months before a rating downgrade. Adjustments in the rating process illustrate a wealth effect as well. Moody's changed its rating categories in April 1982 seeing equity and bond prices to react with the introduction of numeric modifiers (Kliger and Sarig, 2000).

Rating announcements are usually preceded by the same rating agency or initiated by another rating agency. (Norden and Weber, 2004) conducted a study analyzing reviews for downgrades with significant abnormal returns for equity and bond prices but actual downgrades with no significant findings. (Hull et al., 2004) investigated rating announcements from Moody's with similar results. To measure the impact of preceded rating announcements, (Norden and Weber, 2004) find significant results if no rating announcement was done in the last 12 months prior to the event day.

Data and Methodology

Data

We extract data from the following sources. For stock market prices we use Thomson Reuters and for rating announcements Bloomberg, S&P, Fitch and Moody's. We narrow the search for rating announcements by selecting CDO as issue type. In order to link the issuer of the CDO deal to the ultimate parent institution, we need to identify the SPV as the investment vehicle of the issuing bank.

According to the collateral type, there are big differences in the number of tranches. In concordance with the Basel II regulatory framework, rating agencies treat each tranche of a CDO deal independently. Nonetheless, rating agencies can issue rating announcements on several tranches or even several CDO deals on a single day. We prove whether there have been confounding events as these may bias our findings. These confounding events are handled as followed. If rating announcements are published for more than one tranche of a particular deal we do not consider this as a confounding event. We follow this rule as it is common practice of rating agencies to undergo the rating process of more than one CDO tranche of the same deal. Rather do we consider confounding events when there are rating announcements on tranches of more than one CDO deal issued by the same bank in our largest event window [-20; 20]. With this approach we considerably reduce our sample size.

Our research sample consists of CDO from issuing banks headquartered in Germany, Switzerland, UK and USA. At first, we select all rating announcements – outlooks and downgrades in the period from 1999 until June 2010. From 6002 rating announcements for CDO of US banks and 2007 rating announcements from German, Swiss and UK banks, we consolidate the list by filtering CDO issued by stock-listed banks with considerable securitization activity. It is worth mentioning that a great share of CDO is issued through banks that are not listed on the stock market. This leaves us with 1227 rating announcements for US banks and 1743 rating announcements for banks headquartered in Germany, Switzerland and UK. Further on, if there is more than one rating for several tranches of the same CDO deal, we consider this as one event. After that we identify confounding events. We define confounding events not only as rating announcements in our largest event window but disallow major events that are published during the event window in order to prevent analysis bias with corporate news from Financial Times and Wall Street Journal. Exemplary, on the day Lehman Brothers claimed for bankruptcy, rating downgrades of CDOs issued by Lehman Brothers were noticed. As this is a major confounding event, we exclude such an event from our sample.

Our final sample size consists of 72 events from US banks and 167 events from German, Swiss and UK banks. CDOs started to receive investor attention in the past decade. We take this into account and limit our time horizon from 1999 until June 2010. We define a subsample from 1999 until 2006 and a subsample from 2007 until June 2010. Another division of our subsample refers to the region of the issuing bank, Europe or the USA.

Methodology

We measure stock price reactions implicated by rating announcements. For that purpose we apply an event-study approach that is designed to quantify abnormal returns within a specified event period. We measure abnormal returns considering stock market effects with the market model approach as outlined by (MacKinlay, 1997).

We consider market-model adjustments to calculate the cumulative abnormal return (CAR):

$$CAR_{i,[t_0-\tau,t_0+\tau]} = \sum_{t=t_o-\tau}^{t=t_0+\tau} (R_{i,t} - \hat{\alpha}_i - \hat{\beta}_i R_{m,t}) \tag{1}$$

where Ri,t is the return of the issuing bank at time t, Rm,t is the security's market return at time t, α_i and β_i and are parameters derived from the ordinary least square regression with the estimation period beginning 200 days before the event with a lag time of 30 days. The largest event window in our observation is set equal to 41 days, starting 20 business days before and ending 20 business days after a rating announcement. We subdivide the event window into 9 time intervals to control for anticipation and post-announcement effects. The selection of the event windows is repeated for all subsamples.

We use adjusted stock returns that take relevant changes such as dividend payments or stock splits into account. We apply the test statistics of (Boehmer et al., 1991) to test the significance of cumulative abnormal returns (CAR) as suggested by (Harrington and Shrider, 2007).

Results

Stock Price Reactions to Rating Announcements

In a first analysis we investigate whether the market considers rating announcements of CDO deals independently from the issuing institution. The results in the overall 239 event sample indicate that highly significant abnormal returns with CAR= -0.96% (t-value=2.604) have occurred right on the event day [0;+0]. The results suggest that CDO deals are not independently considered from the originating bank and we are going to challenge this assumption. Our sample shows furthermore that rating announcements are absorbed by market participants very quickly. In the event windows [-1;+2], [-1;+1] and [0;+2] we get significant abnormal returns at the 5% confidence level. This perception is in line with the corresponding finance literature regarding announcement returns to debt downgrades (Higgins et al., 2009). In the introductory part we raised the question if the implicit recourse, a characteristic of CDO deals, does not fully transfer the risk from the issuing bank to the SPV as losses of the reference portfolio strikes the first loss piece of the originating bank. The results table outlines that risk still resides with the issuing bank and so is not fully transferred. With this in mind the true sale assumption, meaning an effective risk transfer, is violated. On a longer period of event windows we get different results. Once the rating announcements are absorbed, there are no abnormal returns noticed in larger event windows [-5;+5] or [-10;+10]. Interestingly, the mean and median of the largest event window [-20;+20] deliver positive values. With a total sample of 239 events it appears that shareholders of the issuing bank do not consider CDO rating announcement as a prime driver in their investment decisions.

The table points out that there are considerable differences in the mean and median values assuming that some negative abnormal returns have a strong weight on the mean value. Therefore we conduct this analysis with value weighted results and obtain similar results. We mentioned that the growing importance of rating agencies could force investors to blindly rely on rating grades as a quality measurement. Due to the in-depth view on CDO deals and experienced staff in analyzing CDOs, the significance of rating agencies could be wrongfully overrated.

We challenge whether rating announcements have a wealth effect independently from regional focus or time periods. Two options are going to be discussed. Does the market either comport itself indifferent to rating announcements, what violates the assumption of the overrated impact of rating grades, or does it anticipate upcoming rating announcements. We neglect the first assumption of indifferent market reaction as the overall result table indicates significant wealth effects around the announcement day. The letter assumption referring to anticipation effects is of key interest in our remaining analysis.

Table 1. Parent Institution Market Reactions: Market Model - Estimation Period: [t-200;t-30] - equally weighted, 239 events from 1999-2010

	Cumulative Abnormal Return		***t-Test***	***Boehmer Test***	***Wilcoxon signed rank test***	
Event window	***Mean***	***Median***	***t-value***	***z-score***	***z-score***	***Nobs***
[-10;+1]	-1.33%	-0.39%	-1.279	-1.436	-0.940	239
[-5;+1]	-1.17%	-0.33%	-1.228	-1.304	-1.036	239
[-1;+1]	-1.18%	-0.22%	-2.281**	-2.087**	-3.181***	239
[-5;+5]	-0.48%	-0.22%	-0.903	-0.997	-1.169	239
[-10;+10]	-0.66%	-0.15%	-0.905	-1.030	-0.575	239
[-20;+20]	0.30%	0.02%	-0.260	0.361	-0.064	239
[0;+0]	-0.96%	-0.25%	-2.604***	-2.351**	-4.646***	239
[-1;+2]	-1.13%	-0.44%	-2.193**	-2.036**	-2.929***	239
[0;+2]	-1.01%	-0.29%	-2.080**	-1.976**	-2.756***	239

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

In the introductory part we mentioned that first market distortions in the subprime credit business were noticed in 2006. Subprime mortgages have accounted for a great share of the reference assets in CDO deals. If the investors were aware of the securitization and CDO activity of US banks, they might have anticipated the upcoming downgrade wave for CDO acting indifferent to rating announcements. In a subsample concerning 72 rating events corresponding to US banks we confirm the anticipated rating announcements.

Table 1. Parent Institution Market Reactions: Market Model - Estimation Period: [t-230;t-30] - equally weighted, 72 events for US banks

	Cumulative Abnormal Return		***t-Test***	***Boehmer Test***	***Wilcoxon signed rank test***	
Event window	***Mean***	***Median***	***t-value***	***z-score***	***z-score***	***Nobs***
[-10;+1]	-4.40%	-1.11%	-1.569	-1.559	-2.233**	72
[-5;+1]	-3.56%	-0.73%	-1.268	-1.252	-1.235	72
[-1;+1]	-2.70%	-0.71%	-1.610	-1.492	-3.109***	72
[-5;+5]	-0.72%	-0.32%	-0.766	-0.678	-0.617	72
[-10;+10]	-2.04%	-0.74%	-1.589	-1.548	-1.341	72
[-20;+20]	1.18%	-1.08%	-0.257	0.628	-0.690	72
[0;+0]	-1.97%	-0.22%	-1.533	-1.493	-2.581***	72
[-1;+2]	-2.41%	-0.62%	-1.483	-1.383	-2.699***	72
[0;+2]	-2.04%	-0.39%	-1.314	-1.272	-1.678*	72

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

All of the event windows do not illustrate significant results and notify that rating announcements were anticipated by the market well in advance of the event day. From our overall result table we have seen that there are significant results around the announcement day. So we investigate wealth effects of European banks issuing CDOs. In the findings of (Uhde and Michalak, 2010), securitizations have increased systematic risk of European banks. This may result in negative wealth effects for European banks.

In line with the overall sample, wealth effects are most pronounced around the announcement day. Parametric and non-parametric tests show highly significant negative abnormal returns. In the event window [0;+0] we obtain negative abnormal returns of CAR=-0.52% (t-value=-3.906). This indicates, in contrary to CDO from US banks, that the market did not anticipate the rating announcements. The result table supports our findings from the overall result that significant abnormal returns occur quickly around the announcement day in our shortest event windows.

Table 2. Parent Institution Market Reactions: Market Model - Estimation Period: [t-200;t-30] - equally weighted, 167 events for European banks

	Cumulative Abnormal Return		***t-Test***	***Boehmer Test***	***Wilcoxon signed rank test***	
Event window	***Mean***	***Median***	***t-value***	***z-score***	***z-score***	***Nobs***
[-10;+1]	-0.01%	-0.16%	0.152	-0.013	-0.221	167
[-5;+1]	-0.14%	-0.14%	-0.164	-0.376	-0.463	167
[-1;+1]	-0.53%	-0.16%	-2.157**	-2.416**	-1.795*	167
[-5;+5]	-0.38%	-0.18%	-0.571	-0.728	-0.904	167
[-10;+10]	-0.07%	0.05%	-0.065	-0.095	-0.198	167
[-20;+20]	-0.09%	0.99%	-0.138	-0.100	-0.505	167
[0;+0]	-0.52%	-0.29%	-3.906***	-4.063***	-3.842***	167
[-1;+2]	-0.58%	-0.27%	-2.037**	-2.237**	-1.763*	167
[0;+2]	-0.57%	-0.29%	-2.169**	-2.368**	-2.216**	167

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

To further analyze whether market participants rely on rating grades we pursue the analysis as follows. CDO is a rather young investment opportunity and therefore we investigate the effect of time dependency. Over the last decade CDO have gained market attention due to beneficial facts such as capital relief and credit risk transfer. We split our analysis in a time period from 1999 until 2006 and a time period from 2007 until June 2010.

As first market distortions in the credit business have been noticed in 2006 and the awareness of high default rates became apparent, investors should be surprised before the credit crunch. Once the weak performance of many subprime credits came to public attention, investors should not be surprised when the actual announcement has been disclosed. This hypothesis is in line with our earlier findings that rating announcements of CDOs from US banks – most rating announcements took place after 2007 – were anticipated.

Table 3. Parent Institution Market Reactions: Market Model - Estimation Period: [t-200;t-30] - equally weighted, 127 events from 1999-2006

	Cumulative Abnormal Return		***t-Test***	***Boehmer Test***	***Wilcoxon signed rank test***	
Event window	***Mean***	***Median***	***t-value***	***z-score***	***z-score***	***Nobs***
[-10;+1]	0.08%	-0.29%	0.066	0.247	-0.233	127
[-5;+1]	0.02%	-0.14%	0.021	0.085	-0.472	127
[-1;+1]	-0.41%	-0.34%	-2.674***	-2.522**	-2.616***	127
[-5;+5]	0.20%	0.29%	0.592	0.652	-0.440	127
[-10;+10]	0.52%	-0.19%	0.848	1.152	-0.479	127
[-20;+20]	0.36%	0.44%	0.391	0.536	-0.575	127
[0;+0]	-0.27%	-0.18%	-2.772***	-3.029***	-2.644***	127
[-1;+2]	-0.28%	-0.36%	-1.547	-1.443	-1.684*	127
[0;+2]	-0.11%	-0.17%	-0.716	-0.744	-0.946	127

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

In the first time period from 1999 until 2006 wealth effects are noticed with CAR= -0.27% (t-value=-2.772) at announcement day. For the event period [-1;1] the CAR is highly significant as well. In accordance to our findings from the overall sample and CDO issued by European banks we do not receive significant results in our longer event periods and confirm that the rating information is quickly incorporated.

Rating announcements from 2007 until June 2010 show significant results primarily on the announcement day that stands in contradiction of anticipated rating events. The evidence cannot be stressed as strongly as in the earlier findings due to the lower significance level. The parametric test indicates significant results at the 5% level.

Table 4. Parent Institution Market Reactions: Market Model - Estimation Period: [t-200;t-30] - equally weighted, 111 events from 2007

	Cumulative Abnormal Return		***t-Test***	***Boehmer Test***	***Wilcoxon signed rank test***	
Event window	***Mean***	***Median***	***t-value***	***z-score***	***z-score***	***Nobs***
[-10;+1]	-2.92%	-0.42%	-1.355	-1.494	-1.003	111
[-5;+1]	-2.50%	-0.48%	-1.237	-1.308	-0.900	111
[-1;+1]	-2.09%	-0.22%	-1.652*	-1.734*	-2.113**	111
[-5;+5]	-1.15%	-0.76%	-1.242	-1.174	-1.615	111
[-10;+10]	-1.88%	0.03%	-1.489	-1.477	-1.015	111
[-20;+20]	0.42%	-0.40%	-0.414	0.265	-0.235	111
[0;+0]	-1.75%	-0.44%	-2.069**	-2.015**	-3.890***	111
[-1;+2]	-2.10%	-0.78%	-1.782*	-1.792*	-2.419**	111
[0;+2]	-2.03%	-0.58%	-1.945*	-1.880*	-2.757***	111

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Anticipation Effect of Rating Announcements

We mentioned in our literature review that some empirical studies found evidence that upcoming rating announcements were anticipated by market participants. Significant negative abnormal returns for bonds 90 days prior to a rating downgrade or negative watch listings was found by (Steiner and Heinke, 2001). Contrary results were found by (Weinstein, 1977) with no abnormal performance 6 months before and after the event. As we analyze the implication of bond downgrades on the issuing bank's equity we rather focus on the findings of (Hand et al., 1992). They found evidence of significant negative abnormal stock and bond returns for downgrades. (Hull et al., 2004) found support that negative rating downgrades were preceded by rising CDS spreads.

In order to conduct our analysis with anticipation and post announcement effects we provide a buy and hold abnormal return (BHAR) long term event study according to (Kothari and Warner, 1997; Lyon et al., 1999). We further consider equally and value-weighted results. In order to account for anticipation effects we limit the pre-event period to 3 months prior to the event. (Pinches and Singleton, 1978) analyzed abnormal returns for stocks 30 months prior to the event time. While this may seem justified for stocks, we only use a short period of 3 months as issuing CDO deals is not a core business of a bank. The selection of the peer group is essential in long term analysis. (Higgins et al., 2009) propose matching with the industry (4 digit SIC code), size (market value equity) and book-to-market equity. We select the market index rather than a particular peer bank because most of the stock listed peer banks are engaged in securitization and CDO activity. Therefore, the selection of peer banks might bias our results.

If rating downgrades are expected and already priced, we would expect that significant abnormal returns take place right before the actual event. The long term effect for CDO is slightly different. Market participants were well aware of the upcoming wave of rating downgrades for structured finance product, particularly CDO. Referring to our above mentioned note we pointed out that CDO is not a core business of a bank. Hence, we would expect that long term significant abnormal returns were incorporated in the share price of the issuing bank well before the event date and meanwhile the share price movement is not primarily driven by rating announcement of CDOs. So, pre event BHAR should not illustrate significant abnormal returns.

Table 5. Pre-announcement BHAR

BHAR		*t-Test*	*Johnson Test*	*Wilcoxon signed rank test*	
Mean	***Median***	***t-value***	***J-value***	***z-score***	***Nobs***
-0.25%	0.41%	-0.287	-0.287	-0.050	239
0.19%	-0.19%	0.184	0.185	-0.035	239
0.17%	-0.19%	0.128	0.129	-0.539	239

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Post Event Effect of Rating Announcements

Post event BHAR reveal that there are no long term effects of CDO downgrades on the issuing bank's share price. We used as the maximum observation time three months after the event day. Equally weighted results do not show significant abnormal returns. This supports our findings from the short term event study that in our largest event window [-20; 20] downgrades of CDOs were not substantial to the issuing bank's share price. Hence, market participants were, especially during the financial turmoil, well aware of upcoming rating downgrades and were not surprised when the actual downgrade took place on a longer period.

Table 6. Post-announcement BHAR

BHAR		*t-Test*	*Johnson Test*	*Wilcoxon signed rank test*	
Mean	***Median***	***t-value***	***J-value***	***z-score***	***Nobs***
-0.72%	-0.31%	-0.763	-0.761	-0.807	239
-0.41%	0.02%	-0.337	-0.335	-0.345	239
0.40%	-0.53%	0.283	0.283	-0.064	239

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Multivariate Regression Analysis

This section reports regression analysis of the issuing bank's returns on explanatory variables designed to identify the drivers of abnormal returns in section 4.1. The dependent variable is the CAR from the event window. The independent variables test the impact on the CAR in the shortest event window [0;+0] by using deal characteristics. According to our subsamples presented in section 4.1 we extend our analysis by introducing dummy variables equal to one if the issuing bank is from Europe (D_{eur}) and another dummy variable if the rating announcements occurred before 2006 ($D_{before2006}$). As the banking industry in general was in a stress scenario we will consider risk factors in our multivariate analysis. We measure the systematic risk until 2006 in comparison to the systematic risk after 2007. Another important factor is whether the market could assess the sound condition of a bank. We introduce a dummy variable if rating announcements have been disclosed for the issuer in advance to the event day ($D_{issuer\ downgrade}$). (Higgins et al., 2009) examined the ability of the market to assess the risk of a bank when investigating wealth effects of rating announcements on ABS. Negative wealth effects were found for banks that were downgraded prior to the actual downgrade of the ABS. We argue that a core driver of the abnormal returns obtained in section 4.1 goes back to preceded rating downgrade for the issuer.

The t-statistics of the coefficient estimates are based on White's heteroskedasticity-consistent standard errors and are reported in parentheses (White, 1980).

Table 7. Multivariate regression analysis

Multivariate Regression	Coefficients	*t-values*
Intercept	-0.0401	*-3.4184*
Dummy=1, EU issuing banks	0.0016	*0.4983*
Dummy=1, rating announcement until Dec 2006	0.0037	*1.8689*
beta 1999-2006	-0.0010	*-0.8018*
beta 2007-2010	-0.0004	*-1.3122*
Dummy=1, Issuer Rating Downgrade 3 months prior to the rating event	-0.0108	*-2.1349*
Size	0.0077	*3.1914*
Adj. R^2	0.1353	

The results show that there is a general negative impact of rating announcements for CDO with the intercept coefficient -0.0401 (t-value= -3.4148). We see significant negative impact if the issuer experienced a rating downgrade three months before the event day. The systematic risk shift was not a primary driver of the abnormal returns (CAR).

Conclusions

CDO have gained attention over the last years. Regulatory legislation has been favorable for CDO. High yielding CDO deals with investment grade status were demanded as a promising investment opportunity. Rating agencies recognized that providing rating grades for structured credit products like CDO became a considerable profit contributor. Nonetheless, neither the rating agencies nor the market participants were cautious enough to assess the risk CDO were bearing. Beginning with the subprime credit crisis, the high rating grades were questioned by the market seeing a rapid drop in confidence to complex products such as CDO. In the wake of the financial crisis CDO were accused to bear incalculable risk. Risk transfer to a SPV did not fully transfer risk from the originator due to implicit recourse techniques. As this became apparent, investors have forgone banks that were actively involved in CDO issuing. Once the CDO market dried out, market participants did not rely on rating grades anymore. Our sample and results indicate that rating downgrades were anticipated well in advance but were surprising when the actual rating announcement took place. Our subsamples reveal that market mistrust was existent especially for CDO issued by US banks. The underlying assets of European CDO were assumed to be of better quality than CDO from US banks. Still, significant abnormal returns were noticed on the announcement day and quickly incorporated for CDO issued by European banks. As CDO is not a core business of a bank, rather a risk management instrument, with no long term abnormal returns. Overall, we deliver evidence that the rating announcements are overrated once the market participants are aware of upcoming downgrades.

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CEO Turnover and Compensation: Evidence of Labor Market Adjustments

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Abstract

CEO turnovers represent an opportunity for boards of directors to negotiate anew regarding compensation. If compensation has become suboptimal, the board can use this opportunity to resolve any issues that may have risen during the tenure of an outgoing CEO. This paper investigates whether boards take advantage of this opportunity. Specifically, I examine changes in compensation packages when a turnover occurs and relate the changes to conditions existing prior to the turnover. The evidence suggests that boards do make changes to compensation when turnovers occur. Specifically, when outgoing CEOs appear to have been overpaid (underpaid), incoming CEO pay is lower (higher).

Introduction

CEO turnover events provide a special occasion for the restructuring of CEO compensation. Existing literature continues to debate whether CEOs are paid efficiently or their compensation reflects systematic sub-optimality. If compensation has become suboptimal, the board can use this opportunity to resolve any issues that may have risen during the tenure of an outgoing CEO. This paper investigates whether boards take advantage of the opportunity provided by a turnover event. Specifically, I examine CEO compensation levels and "excess pay" (measured as the residual from a standard model of compensation) for a sample of 1,232 incoming and outgoing CEOs involved in turnover events in U.S. public companies between 1993 and 2006.

There are essentially two ways in which CEO compensation can become sub-optimal over the tenure of a CEO. First, the CEO's pay may become inflated and represent some rent extraction in the form of "overpay". This "overpay" is often described in the literature as a function of the CEO's power over the board. That is, since the CEO is responsible for nominating board members and since he has established relationships with board members during the course of his tenure, the "managerial power" view of CEO compensation contends that CEOs both have and exercise power over board members to induce them to pay him more than he is worth (see, for example, Morck, Shleifer, and Vishny (1988), Crystal (1991), Jensen (1993), Bebchuk, Fried, and Walker (2000); and Mullainathan (2001), Garvey and Milbourn (2004); Brenner, Sundaram, and Yermack (2000), Chance, Kumar, and Todd (2000), Pollock, Fisher, and Wade (2002)). When this is the case, shareholders should expect the board to be vigilant and attempt to remedy the situation whenever they can. Second, the CEO's pay may have stagnated and fallen below expectations (based on comparable firms) towards the end of his tenure. If this is the case, shareholders should expect the board to remedy this situation and "catch up" with the managerial labor market when a turnover occurs.

This paper examines the following questions. First, I ask whether the board appears to make corrections in determining an incoming CEO's compensation when it appears that outgoing CEOs have been overpaid or underpaid. Second, I ask whether there is a relationship between "managerial power" and changes in compensation when a turnover occurs. To address the first question, we must first establish a definition of "overpay" and "underpay". I create a measure consistent with prior studies called "excess pay" and consider CEOs to be overpaid if the measure is positive and underpaid if the measure is negative[1]. Specifically, I follow the approach used by Hartzell, Ofek, and Yermack (2004) and measure "excess pay" as the residual from a standard model of CEO total compensation as a function of firm size, firm performance (3-year stock price performance), industry, and year.

Using this measure, I find that approximately 35% of outgoing CEOs are overpaid. Within this set of outgoing CEOs, the median "excess pay" is approximately $2.15 million. In 43% of these cases the replacement does not receive "excess pay" > 0. Thus, it appears to be the case that boards are correcting, but not nearly as much as one might expect if "excess pay" is truly a measure of overpay. However, within the entire sample of outgoing CEOs whose "excess pay" > 0, there is generally and systematically a decrease in pay from outgoing to incoming CEO. Specifically, the incoming CEO total compensation within this subset is, at the mean (median), $2.669 million ($275,269) less than his predecessor's. Cash compensation is $179,565 ($47,279) lower and option awards are worth $2.24 million ($163,340) higher. These differences are all statistically significant at the 1% level. Thus, even when the incoming CEO appears to be overpaid as well, the evidence still suggests some correction. Multivariate analysis strengthens the result that when outgoing CEOs appear to be overpaid, their replacements are paid less.

As a potential explanation, it may be the case that some of the outgoing managers have become entrenched. If so, the expectation should be that when a turnover occurs, any overpay that was the result of managerial entrenchment would be eliminated by the board of directors. To explore this possibility, I examine the relationship between governance measures and changes in pay from outgoing to incoming CEO. First, I examine anti-takeover measures, which are commonly used as a measure of managerial entrenchment and/or poor governance throughout the literature. I use the governance index introduced in Gompers, Ishii, and Metrick (2003), though findings are robust to alternative specifications (i.e. the entrenchment index used by Bebchuk, Cohen, and Ferrell, 2009). I find that within the set of turnovers for which outgoing CEOs appear to be overpaid and boards appear to correct (i.e. outgoing CEO "excess pay"> 0, incoming CEO "excess pay"≤0), the governance index (gindex) is statistically significantly higher by approximately 1. Furthermore, I find that a high governance index (indicating higher levels of entrenchment) is associated with larger decreases in pay from incoming to outgoing CEO and that this effect is especially pronounced when the outgoing CEO appears to have been overpaid. That is, when the gindex is high and the outgoing CEO appears to have been overpaid, the mean (median) pay for the incoming CEO is $4.92 million ($812,000) less than his predecessor. However, when the gindex is average or lower and the outgoing CEO appears to have been overpaid, the incoming CEO receives $1.46 million ($451,000) more than his predecessor. These differences are all statistically significant at the 1% level. Furthermore, logistic regressions reveal that, although the likelihood of the incoming CEO receiving "excess pay" > 0 is higher when the outgoing CEO received "excess pay", the effect is reduced when the firms have a high gindex.

In addition to antitakeover measures, I also examine the impact of outgoing CEO founder status, outgoing CEO ownership, outgoing CEO tenure, board independence, and incoming CEO insider/outsider status. Results are mixed. With respect to managerial power, we would expect CEO founder status, ownership, and tenure to be positively related to CEO power while board independence and outsider status should be negatively related to CEO power. Some results support this. For instance, when the outgoing CEO appears to have been overpaid and had high ownership levels, the pay differential is very large and in favor of the outgoing CEO. However, I find similar results when the outgoing CEO appears to have been overpaid and board independence is higher than average. In multivariate regressions, the interaction between outgoing CEO "excess pay" and high CEO ownership and high board independence is negative and significant. Furthermore, I find that the likelihood of the incoming CEO receiving "excess pay" is negatively related to interaction terms between outgoing CEO "excess pay" and 1) high outgoing CEO ownership and 2) high board independence.

In contrast to the "overpay" cases, over 52% of the entire sample represents turnovers in which neither the incoming nor the outgoing CEO receive "excess pay" > 0 and when the outgoing CEO does not receive "excess pay" > 0, only 16% of incoming CEOs receive "excess pay" > 0. The median "excess pay" within the set of CEOs paid as expected or underpaid is approximately $2.6 million. When the outgoing CEO appears to be underpaid, his replacement receives, at the median, $181,000 more in option-based pay and $490,000 more in total compensation. These figures are also statistically significant at the 1% level. His cash-based pay is also higher and statistically significant, however the median value is only $33,000. Thus, while the "excess pay" measure indicates below-expected pay for the incoming CEO, it is significantly higher than his predecessor's and thus may reflect a board's attempt to "catch up" with the managerial labor market.

As yet, I am aware of limited research that has explicitly and comprehensively examined compensation contracts around turnover events. Rather, most papers examine turnover in the context of its being an individual potential component of the overall governance package that is designed to align managers' incentives and curb rent extraction – that is, the threat of turnover is typically viewed as playing a disciplinary role in the context of the manager-shareholder agency problem. These studies typically examine the occurrence of management turnover in relation to firm performance and other firm or market characteristics (see, for example, Denis, Denis, and Sarin (1997); Coughlan and Schmidt (1985); Murphy and Zimmerman (1993); Huson, Parrino, and Starks (2001); Goyal and Park (2001); Lehn and Zhao (2006)). The main contribution of this paper is two-fold: first, I examine the dynamic interaction between two governance mechanisms (turnover and compensation) and second, I examine potential measures of CEO overpay/underpay.

In spirit, this paper resembles a varied set of studies regarding CEO compensation. Gilson and Vetsuypens (1990) examine the nature of compensation packages for financially distressed firms and include a discussion of changes observed when a turnover event occurs within this set of firms. They find that, within a small sample of financially distressed firms, when a turnover occurs, insider replacement CEOs are paid substantially less than their predecessors, but outsider replacement CEOs are paid substantially more. Schwab and Thomas (2005) closely examine the negotiation of and provisions in CEO employment contracts, focusing on the legal characteristics of the contract. Murphy (2002) compares levels of CEO pay for insider and outsider replacement CEOs, finding that outsider replacements are typically paid more than insider replacements. Murphy and Zabojnik (2007) expand upon this, presenting a market-based model for compensation that portrays rises in CEO pay, higher pay for outsider replacement CEOs, and the increased prevalence of the appointments of outsider replacement CEOs all as consequences of the need for general managerial skills (rather than firm-specific managerial ability). They document indirect evidence of a shift in the need for general managerial skills and higher pay for outsider replacements.

This study expands upon Murphy (2002) by examining differences between the incoming and outgoing CEOs in turnover events, rather than differences between replacement CEOs' pay based on whether they are insiders or outsiders. Blackwell, Dudney, and Farrell (2007) also expand upon Murphy (2002) and do examine changes in compensation structure following turnover events and relate those changes to firm performance. They find that incoming CEOs' compensation is comprised of significantly more equity-based pay and a positive association between post-turnover performance and new stock grants. Elsaid and Davidson (2009) perform a study very similar to this one. They examine differences between incoming and outgoing CEO pay and the percentages that salary and "pay-at-risk" (i.e. stock and option grants) contribute to total compensation surrounding turnover events.

This paper complements and extends these papers along several lines. First, I examine the dynamic associated with whether the outgoing CEO appears to have been underpaid or overpaid prior to the turnover. To my knowledge, this is the first attempt to determine whether prior potential overpay or underpay has an impact on board compensation decisions at the time of a turnover. Second, I examine a much more comprehensive set of potential CEO compensation determinants and attempt to differentiate between CEO, turnover, and firm characteristics that may contribute to managerial power in a positive way and those that may contribute to some types of "inappropriate" managerial power (i.e. power over the board based on relationships and/or entrenchment effects). Lastly, a large part of the focus in both papers is on changes in the structures of compensation packages. This paper extends the examination to changes in levels of pay and relates them to relative managerial power.

The remainder of this paper is organized as follows: Section II describes the sample and presents preliminary results; Section III contains additional analysis regarding what happens to CEO compensation when a turnover event occurs; and Section IV concludes.

Sample Selection and Summary Statistics

The initial sample is collected from the Standard & Poor's ExecuComp database, which provides information on firms in the S&P 500, the Midcap 400, and the Smallcap 600, between 1993 and 2006. Data collected includes CEOs' cash pay, total compensation including the value of option grants and other forms of pay, tenure as CEO, and CEO percent equity ownership in the firm. ExecuComp provides information regarding the years during which an executive becomes the CEO and leaves office as CEO. Outgoing CEOs are identified by the year in which they leave office as CEO. Incoming CEOs are identified by the year in which they become CEO. The sample of turnovers is then constructed by matching outgoing and incoming CEOs on firm and year. The sample is then limited to those observations for which salary, bonus, and total compensation data is available for the last full year of pay of the outgoing CEO and the first full year of pay of the incoming CEO. Firm characteristics are retrieved from the Compustat database and board and governance data collected from their respective IRRC databases. The sample is reduced to those cases in which at least full firm characteristic data and compensation data are available. This results in 1,232 turnover events over the 13 year period.

Data regarding the insider/outsider status for replacement CEOs specifically is collected from Bloomberg People Search. Bloomberg People Search provides profiles including career history. CEOs are classified as insiders if they have a prior employment history with the firm in which they become CEO. Specifically, where available, if the Bloomberg People Search career history indicates employment with the firm in positions other than CEO prior to the appointment, they are considered insiders. Otherwise, he is considered an outsider. Where Bloomberg People Search career histories are unavailable, Execucomp data is used to determine insider/outsider status. This occurs in few cases, but represents approximately 5% of the sample of incoming CEOs. In this case, CEOs are considered insiders if 1) the first year of the CEO's employment at the firm (according to ExecuComp) differs from the year in which he becomes CEO and 2) the executive appears in the ExecuComp database for that firm for the year prior to the year in which he becomes CEO. The final insider/outsider designation yields a sample consisting of approximately 32% outsider replacements. This is consistent with the relevant prior literature. Agrawal, Knoeber, and Tsoulouhas (2006) use a sample consisting of 18% outsider replacements, but cover a much earlier time period (1974-1995). Murphy and Zabojnik (2009) show that the prevalence of hiring outsider replacements has increased over time, specifically noting that by the year 2005, outsider replacements account for roughly 40% of all CEO replacements.

Founder status is identified via a combination of Bloomberg People Search, news searches, online company histories, and company web-sites. Specifically, when Bloomberg specifies that a CEO is a founder, this is used. If the CEO is not identified as a founder, the news search regarding the turnover event is double-checked for any background information on the individuals. This allows for a designation of both founder status and family status. If founder status is still not found through either of these sources, I search for company histories online using Google. If founder status is still not found, I go directly to the company's web-site. Board size and composition (measured as the percentage of directors who are independent) are obtained through the RiskMetrics Directors database, which provides detailed data regarding board members of a large number of firms. The governance index first introduced by Gompers, Ishii, and Metrick (2003), commonly referred to as simply the gindex, is acquired through the RiskMetrics Governance database, which provides data regarding various governance characteristics of firms, including the constructed gindex.

Using an approach similar to Hartzell, Ofek, and Yermack (2004), "excess pay" is calculated as the residual from a basic regression model. That is, using the entire universe of ExecuComp data, I run the following regression:

$$\text{CEO Pay} = \log(\text{MktCap}) + \text{3-year stock returns} + \textit{Industry Dummies} + \textit{Year Dummies} \tag{1}$$

CEO Pay represents total compensation including restricted stock, payouts from long-term plans, benefits, and stock options valued at the grant-date using ExecuComp's modified Black-Scholes methodology. Log(MktCap) is the log of number of shares outstanding multiplied by the year-end stock price. The residuals for the sample of turnover CEOs are retained and used as an estimate of "excess pay". I then use the sign of "excess pay" as a proxy for apparent CEO overpay or underpa[3].

Overall, mean (median) CEO cash pay and total pay are, respectively, $1.36 million ($957,526) and $4.39 million ($2.24 million). Median CEO ownership of the firm for outgoing CEOs is 1.4%, almost double the median for their replacements of .76%. Median positive "excess pay" is approximately $2.15 million. Median negative "excess pay" is approximately $2.6 million. Insiders account for 68% of replacement CEOs.

On average, the sample represents large firms with mean (median) assets of close to $13 billion ($2 billion), mean (median) sales of $5.5 billion ($1.52 billion), mean (median) net income of $102 million ($50 million), and mean (median) total debt of $4.3 billion ($435 million). Median return on assets is 3.73% and the median prior 3-year stockholder returns is 7.38%. Average board sizes are around 10 board members and the average percentage of independent board members on the board is around 66%. The average gindex for the sample is 9, indicating that the average firm will have 9 antitakeover measures in place.

Firm performance, measured as firm 3-year stock returns, is consistently worse prior to an external hire and in cases in which the outgoing CEO's "excess pay" > 0. This could reflect an entrenchment effect or rent extraction. In particular, the median 3-year stock return for firms which appear to be overpaying their CEOs is 3.2%. In contrast, the analogous value for all other firms is almost 9%. Net income is also significantly lower for firms that exhibit outgoing CEO "excess pay" > 0.

Table 1 shows the sizes of the partitioned subsamples based on insider/outsider designation and "excess pay".

Table 1: "Excess Pay" Subsets

	Incoming CEO Excess Pay ≤ 0	Incoming CEO Excess Pay > 0	Total
Panel A: Excess Pay			
Outgoing CEO Excess Pay ≤ 0	52.45%	12.50%	64.94%
Outgoing CEO Excess Pay > 0	14.85%	20.21%	35.06%
Total	67.30%	32.70%	100%
Panel B: Outsider Replacements			
Outgoing CEO Excess Pay ≤ 0	43.33%	16.19%	59.52%
Outgoing CEO Excess Pay > 0	14.76%	25.71%	40.48%
Total	58.09%	41.90%	100%
Panel C: Insider Replacements			
Outgoing CEO Excess Pay ≤ 0	54.31%	11.74%	66.05%
Outgoing CEO Excess Pay > 0	14.87%	19.08%	33.95%
Total	69.18%	30.82%	100%

There is consistently a set of around 15% of the sample for which the outgoing CEO was paid "in excess" and his replacement is not. This subset may be an important one if evidence can be found to indicate that it represents a set of firms in which the board of directors is attempting to remedy inefficiencies. Alternatively, this subset could simply be reflecting some mean reversion in CEO compensation. This is examined more closely in the subsequent section.

Table 1 also shows that the largest partition is consistently that for which both incoming and outgoing CEOs are not paid "in excess". If a discussion of the two main competing theories regarding compensation is to end with the conclusions that both are at least partially correct, I would expect that this partition is one in which we would find little support for the managerial power theory. On the other hand, the sample for which both CEOs are paid "in excess" may represent those cases in which the managerial power theory is holding true. In a later section, I examine these subsets more closely.

Table 2 presents the differences from paired CEOs surrounding the sample turnover events. The incoming CEO's first full year of pay is subtracted from the outgoing CEO's last full year of pay (I will refer to this as the "pay differential" from this point forward). Univariate analysis is then run on the differences. Overall, median total compensation is $159,560 higher for incoming CEOs than for their predecessors, though median cash pay is not significantly different. The difference in total compensation is significant at the 1 percent level. Consistent with Murphy (2002), Table 2 also shows that outsider replacement CEOs, at the median, typically make $335,360 more than their predecessors while insiders are typically paid only $126,156 more than their predecessors. However, the difference in differences between insider and outsider replacements is not statistically significant. Lastly, Table 2 shows that when the outgoing CEO's "excess pay" ≤ 0, his replacement is much more consistently paid more than he was and vice versa. That is, when the outgoing CEO's "excess pay" ≤ 0, the median replacement CEO is paid $490,343.50 *more* than his predecessor. When the outgoing CEO's "excess pay" > 0, his replacement is paid $275,269 *less* than his predecessor. These are all significant at the one percent level. Furthermore, the differences in the differentials are also significant at the one percent level.

Table 2: Changes in compensation

Median Differences in CEO Pay: Incoming CEO – Outgoing CEO					
	All CEOs	Replacement Insiders	Replacement Outsiders	Outgoing CEO Excess Pay > 0	Outgoing CEO Excess Pay ≤ 0
Total Cash Pay ($)	0	-35,256**	96,557***	-47,279***	33,789*
Total Compensation ($)	159,560***	126,156***	335,360***	-275,269***	490,343***

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

Next, I take a closer look at the interaction between insider/outsider replacements and outgoing CEO excess pay. Table 3 displays pay differentials on this basis. When the outgoing CEO appears to have been overpaid, the median pay differential is negative and significant, regardless of whether the incoming CEO is an outsider. The differences between insider and outsider pay differentials are not significant when outgoing CEO excess pay is positive. This changes, however, when the outgoing CEO does not receive positive excess pay. In this setting, outsiders receive $641,371 more than their predecessors in total pay and $156,448 more in total cash pay. In contrast, insiders only receive $375,277 more than their predecessors in total pay and there is not a significant difference between insider replacements' total cash pay and their predecessors'. The differential for total compensation between insider and outsider replacements is significant at the one percent level. Thus, it appears that the overall results regarding insider/outsider pay differentials are driven entirely by the subset in which outgoing CEOs do not appear to be overpaid.

Table 3: Changes in compensation, Excess Pay and Insider/Outsider Status

	Excess Pay, Insiders	Excess Pay, Outsiders	No Excess Pay, Insiders	No Excess Pay, Outsiders
Total Cash Pay ($)	-65,000**	-22,639	-15,154	156,448***
Total Compensation ($)	-256,110***	-283,230***	375,277***	641,371***

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

Next I turn to the interaction between excess pay and potential sources of managerial entrenchment. For this analysis, I examine the impact of board composition, founder status, CEO tenure, CEO firm ownership, and GIM-index on compensation surrounding the turnover events. These are all considered throughout the literature to be associated with managerial power either directly (as in founder status and tenure) or indirectly via their impact on the quality of governance mechanisms. Donaldson and Lorch (1983), Finkelstein (1992), and Adams, Almeida, and Ferreira (2005) are just a few of the papers that discuss the increased influence of ownership and founder status on the board of directors. Tenure as CEO undoubtedly is positively associated with a CEO's influence over the board. Furthermore, Gregory-Smith, Thompson, and Wright (2009) find that the likelihood of CEO departure drops dramatically after a CEO's fourth year in office and that this appears to be due to CEO entrenchment. Independent (or outside) board members are expected to be farther removed from the CEO (and have a more tenuous relationship with the CEO) and thus are expected to have a negative impact on managerial power. However, Cyert, Kang, and Kumar (1997) find a positive association between CEO compensation and percentage of outside directors on the board. They generally attribute this to increased incentives in the form of option-based pay, which leads to higher levels of total compensation. The GIM index constructed by Gompers, Ishii, and Metrick (2003) has largely been used to measure the quality of corporate governance. It is constructed by summing indicators for various 24 charter provisions, bylaw provisions, and other firm-level rules associated with hostile bidders, voting rights, director/officer protection, other takeover defenses, and state laws. A higher index score represents greater managerial power and weaker shareholder rights.

Tables 4 and 5 show the results regarding changes in compensation broken down by excess pay, tenure, and outgoing CEO ownership. High tenure is defined here as a dummy variable that is equal to 1 when the outgoing CEO has been in office for four years or longer. I use 4 years as a cutoff based on Gregory-Smith, Thompson, and Wright (2009), who find evidence of entrenchment effects and lower probabilities of forced turnovers after the fourth year in office. High ownership is defined as CEO firm ownership being higher than the average for the sample. However, in untabulated results, the analysis is run with higher thresholds (5% and 10% firm ownership[3]). In Table 4, we see that high ownership levels have little impact on pay differentials when the outgoing CEO received positive excess pay. Tenure does appear to have an impact. The differential moves in favor of the incoming CEO when the outgoing CEO has high tenure (from -$369,630 to -$220,700). However, the two differentials are not statistically significantly different from each other.

On the other hand, Table 5 displays results when the outgoing CEO did not receive excess pay > 0. In this case, both high tenure and high ownership result in a higher differential in favor of the replacement CEO. These differentials are statistically significantly different from each other. When the outgoing CEO has high tenure and appears to have been underpaid, his replacement receives $545,442 more than he did. This could indicate some stagnation in pay for those outgoing CEOs with high levels of tenure. This could also be very closely intertwined with high ownership, since we expect that option and stock grants over the course of a CEO's tenure will increase his ownership in the firm (Hambrick, 1995). When the outgoing CEO holds high ownership stakes in the company and appears to have been underpaid, his replacement receives $836,234 more.

Table 4: Changes in compensation: Outgoing CEO Excess Pay > 0, Tenure, and Ownership

	Low Tenure	High Tenure	High Ownership	Low Ownership
Total Cash Pay ($)	-23,822	-59,000**	36,281	-50,221**
Total Compensation ($)	-369,630***	-220,700***	-282,270**	-268,370***

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

Table 5: Changes in compensation: Outgoing CEO Excess Pay ≤ 0, Tenure, and Ownership

	Low Tenure	High Tenure	High Ownership	Low Ownership
Total Cash Pay ($)	161,946***	-863	93,273**	10,500
Total Compensation ($)	372,341***	545,442***	836,723***	438,764***

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

Generally, the results regarding tenure and ownership continue to support the notion that boards will adjust pay in the direction expected when given the opportunity (by a turnover event). However, they do not indicate that outgoing CEO entrenchment plays a role in this context. This is not particularly surprising, given the dichotomous relationship between both CEO tenure and ownership and our expectations regarding their impacts on the firm. On the one hand, each gives the CEO more power over the management of the company (and potentially over the board of directors). However, each is also expected to have a positive impact on the firm. Higher tenure corresponds with more experience running the firm. Higher ownership corresponds with better incentive alignment. Indeed, it is quite possible that compensation levels are lower for outgoing CEOs with high levels of ownership in the firm because the need to align their incentives with those of shareholders is reduced by their ownership levels. Morck, Shleifer, and Vishny (1988), Hermalin and Weisbach (1987), and McConnell and Servaes (1990) demonstrate a non-monotonic relationship between firm value and insider ownership that is believed to reflect the trade-off between the incentives and entrenchment effects of ownership. They all show that concentrated ownership first enhances firm value, but then, at differing thresholds, starts to detract from firm value. Thus, we will not be able to isolate the potential entrenchment impact of ownership and/or tenure in this setting. In untabulated results, I find that founder status has essentially the same qualitative effect as those high ownership.

Next, we turn to board independence and antitakeover measures in an effort to identify any impact of entrenchment on changes in pay when a turnover occurs. Tables 6 and 7 show the changes in pay based on these measures of corporate governance. I define high antitakeover measures as those firms for which the G-index is in the 75^{th} percentile. This equates to those firms with a G-index higher than 11. I define high board independence as those boards which have more than 67% of board members considered to be independent (or outsiders to the firm). The results in Table 6 indicate that, when the outgoing CEO receives positive excess pay, there is a much larger pay decrease (from outgoing CEO to incoming CEO) when the firm has high antitakeover measures and when the board exhibits above-average independent representation. When outgoing CEO excess pay > 0 and the firm has high antitakeover measures (indicating potential managerial entrenchment for the outgoing CEO), the incoming CEO receives \$811,580 (at the median) less than his predecessor. In contrast, when the firm does not have high antitakeover defenses, the differential is only -\$222,140. The differences are even more striking when we consider board independence. High levels of board independence show a remarkable decrease in pay of \$1,425,740 as compared to only \$141,280 when board independence is low. The differences in the differentials are significant at the one percent level.

Table 6: Changes in compensation: Outgoing CEO Excess Pay > 0, Gindex, and Board Independence

	High G-index	Low G-index	High Independence	Low Independence
Total Cash Pay (\$)	-179,134**	-37,714*	-198,000**	-25,467
Total Compensation (\$)	-811,580***	-222,140***	-1,425,740***	-141,280***

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

Table 7: Changes in compensation: Outgoing CEO Excess Pay ≤ 0, Gindex and Board Independence

	High G-index	Low G-index	High Independence	Low Independence
Total Cash Pay (\$)	-74,344	38,359	31,750*	33,789
Total Compensation (\$)	450,610***	493,074***	474,320***	502,668***

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

These results indicate that managerial entrenchment may have an impact on board decisions surrounding a turnover. Specifically, if antitakeover measures (G-index) are high, we expect the outgoing CEO to be more entrenched, all else equal. When high entrenchment is coupled with excess pay, we observe that the board makes larger adjustments regarding the replacements' pay packages. However, Table 7 shows that we observe virtually no difference in differentials when outgoing CEO excess pay is zero or negative. The results regarding board independence, however, indicate that more independent boards are more likely to make larger adjustments when it appears that outgoing CEOs have been overpaid.

Additional Analysis

Table 8 presents the results of multivariate analysis on changes in pay when a turnover occurs. Specifically, I run regressions in which the dependent variable is the pay differential for either total cash pay or total compensation. The dependent variables of interest are insider/outsider status, the measures of governance discussed above, a dummy variable indicating whether the outgoing CEO received positive excess pay, and interaction terms between governance variables and excess pay. To remain consistent with prior literature on CEO compensation, I control for firm characteristics that are expected to have an impact on compensation. These controls include firm size, book to market ratio, prior firm performance (measured as ROA[4]). However, because the dependent variable takes the difference between pay for two CEOs at the same firm, it is not surprising that these firm characteristics do not significantly impact the model. The specific regression equation used is:

Log(Incoming CEO compensation – Outgoing CEO compensation) = Log(Assets) + Book to Market + ROA (2)
*+ Excess Pay Dummy + High Outgoing CEO Tenure + High Outgoign CEO Tenure * Excess Pay Dummy*
*+ Insider + Insider*ROA + High Outgoing CEO Ownership*
*+ High Outgoing CEO Ownership * Excess Pay Dummy*
*+ High G-index + High G-index*Excess Pay Dummy + High Board Independence*
*+High Board Independence * Excess Pay Dummy + e.*

The results are generally consistent with univariate results above, though there are some differences. Outgoing CEO excess pay is highly negatively and significantly related to pay differentials, particularly when examining differentials in total compensation (as exhibited by the -4.285 coefficient on the dummy variable with a p-value < .0001). Tenure is positively related to total compensation pay differentials, but negatively related to cash pay differentials. So, if outgoing CEOs have had longer tenures, we expect incoming CEOs to receive less cash than their predecessors, but more in total compensation. Despite our strong univariate results regarding excess pay and G-index above, the interaction term between the two, while negative, is not statistically significant. However, the interaction term between high ownership and excess pay is negative and highly significant for total compensation differentials. This is a much stronger result than found above in the univariate results. Similarly, the interaction between high board independence and excess pay is highly significant and negative. This is consistent with the univariate results discussed above.

Table 8: Multivariate Regressions: Dependent Variable = Incoming CEO Pay – Outgoing CEO Pay

	Total Cash Pay		Total Compensation	
	Model 1	Model 2	Model 1	Model 2
Intercept	1.245	1.715	1.992*	1.296
Log of Assets	-.00649	.1146	.0303	.1873
Book to Market	-.02081	-.0207	.0253	.0156
Tenure	-.05180**	-.0482*	.0759***	.0554*
Insider	-1.667***	-1.608***	-.4818	-.5984
ROA	-.02125	-.019	.0151	.0111
Insider*ROA	-.02386	-.026	.0582**	-.0567**
Ownership	.1667	-.021	.0575	-.037
High Gindex	-.66067	-.603	-.6607	-.6427
Excess Pay	-1.416***	-1.32*	-4.77***	-4.285***
High Gindex*Excess Pay	-.61578	-.3887	-.8023	-1.087
High Ownership		1.277*		1.213
High Ownership * Excess Pay		-.658		-3.242**
High Tenure		-1.078**		.1005
High Tenure*Excess Pay		.408		1.053
High Board Independence		.3067		-.437
High Board Independence* Excess Pay		-1.108		-2.473***
Adjusted R-Squared	.0416		.1158	
Observations	1082		1089	

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

Finally, I turn to logistic regressions to determine whether the likelihood of the total compensation differential being in favor of the incoming CEO or the likelihood of the New CEO receiving positive excess pay is related to our measures of governance and firm characteristics. Specifically, I create a dummy variable that is equal to one if incoming CEO excess pay > 0 and a dummy variable equal to one if the pay differential between incoming and outgoing CEOs is positive. I then regress these dependent variables on the independent variables discussed above. Table 9 presents the results. Here we find some additional support for our univariate findings regarding the G-index.

Table 9: Logistic Regressions: Likelihood of Incoming CEO Excess Pay or Increase in Pay

	New CEO Excess Pay > 0		New CEO Pay > Old CEO Pay	
	Model 1	Model 2	Model 1	Model 2
Intercept	-.9796***	-1.367***	.8494***	.7151***
Gindex	-.0302	0.0006	-.1583	-.1373
Excess Pay	1.722***	2.033***	-1.1998***	-1.1904***
Gindex*Excess Pay	-.5882**	-.8556**	-.2578	-.2723
Insider	-.2226	-.1395	-.2038	-.2385*
ROA	-.0033	-.00358	.0139	.0147**
Insider*ROA	-.0162*	-.0201**	-.0176	-.019**
High Ownership		.3386		.4638**
High Ownership * Excess Pay		-1.144***		-.5948
High Tenure		-.1465		.1504
High Tenure* Excess Pay		.4531		.3695
High Board Independence		.173		-.0281
High Board Independence* Excess Pay		-1.0635***		-.7609***

* significant at the 1% level , ** significant at the 5% level, *** significant at the 10% level

Specifically, the likelihood of the the new CEO receiving excess pay is positively related to the outgoing CEO receiving excess pay, but the interaction term between outgoing CEO excess pay and gindex is negative. This indicates that when the firm appears to have been overpaying their CEO and has, presumably, high outgoing CEO entrenchment, the relationship between outgoing CEO excess pay and incoming CEO excess pay is much weaker. Thus, this could be interpreted as evidence supporting the notion that the board is making larger corrections when a former CEO has been entrenched *and* overpaid. Using the sign of the pay differential as the dependent variable instead, we find that the likelihood of the new CEO receiving higher pay than his predecessor is negatively related to whether his predecessor received excess pay, as expected.

Conclusion

This study extends the literature regarding corporate governance, CEO turnover, and CEO compensation by examining what changes are made to compensation when a turnover occurs. In particular, I test whether it appears to be the case the boards of directors make adjustments to compensation when it appears to be the case that their CEO compensation package appears represent overpay or underpay. I create a measure of "excess pay" using the residuals from a standard regression model of pay on firm size, performance, industry, and year effects. Using that measure, I show that compensation packages change substantially depending on whether the outgoing CEO appeared to be overpaid (i.e. excess pay > 0) or underpaid. When the outgoing CEO appears to be overpaid, the incoming CEO's compensation is much lower. When the outgoing CEO appears to be underpaid, the incoming CEO's compensation is much higher. Further analysis reveals at least a weak link between measures of corporate governance, excess pay, and adjustments made to compensation when a turnover occurs. Specifically, when firms have a very high level of antitakeover measures (high Gindex) and appear to be overpaying their CEO, there is a much larger reduction in pay from outgoing to incoming CEO pay, suggesting that boards are making quite large adjustments given the opportunity provided by the turnover event. I also find that more independent boards also make much larger adjustments when the outgoing CEO appears to have been overpaid. Overall, the results suggest that if CEO pay appears to become suboptimal over the course of a CEO's tenure, the board makes corrections when a turnover gives them the opportunity.

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Notes

1. Results are robust to using an alternative measure of overpay and underpay based on industry averages.
2. Unreported results are available using industry-adjusted total compensation as an alternative measure of overpay/underpay. The results are all qualitatively similar.
3. Morck et al (1988) find that firm value rises as insider ownership increases up to 5%, then declines until insider ownership reaches 25%, then rises again. Thus, 5% appears to be one appropriate threshold to use. Further, 5% allows for the subsample for high CEO ownership to allow for meaningful statistical analysis (there are 117 outgoing CEOs whose ownership levels are above 5% and 64 CEOs whose ownership levels are above 10%)
4. Or 3 year stockholder returns, in untabulated results.

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Ranking Business Schools by Research Productivity: A Ten-Year Study

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Abstract

We analyze articles published in 10 elite business journals for each of five business disciplines between 2000 and 2009 to investigate the comparative productivity across business disciplines. Our dataset includes 25,997 articles written by more than 15,000 authors that are affiliated with more than 2,000 institutions. We find a significant skew in Ph.D.-school contributions with the top-twenty schools accounting for 44% of all articles. MIT is the top school when ranked by graduates' productivity, while Harvard University tops the ranking when we rank by author affiliation. Additional analysis also indicates the relative importance of foreign schools and non-school contributors to these journals as well as the contributions of individual prolific authors.

Introduction

School ranking continues to be a topic of significance among faculty, students, and administrators. At the heart of the issue is the impact on financial returns and funding. In addition to providing "bragging rights", empirical research indicates that graduates of "top" schools get more employment opportunities and command higher salaries because they are generally considered to be better prepared technically and their research is expected to be superior to that of graduates from lesser-ranked schools. Hence, both administrators and the academic community in general rely heavily on school rankings from various sources.

This paper adds to the growing body of faculty productivity literature by examining the empirical evidence relating to research productivity in 5 business disciplines. We consider cutting-edge research to be vital in faculty productivity rankings so we consider articles only in top-rated journals. We use a list of ten journals[1] for each of five disciplines over a ten-year period (2000 – 2009). We believe that restricting our journal choices to ten journals for each discipline is restrictive enough to ensure that only good-quality articles are included and at the same time is sufficiently wide to cover the primary areas of research in each discipline. The ten-year period provides sufficient time to capture major developments in each discipline and by including articles through 2009, our analysis is up-to-date.

Our three top-rated schools based on graduates' production are Massachusetts Institute of Technology (MIT), Harvard University, and University of Chicago in that order. The overall rankings are significantly different from other rankings in each of the five disciplines and also reveal a number of surprises such as the relative contribution of less well-known programs in US or foreign schools. We also note that almost forty-five percent of articles in our dataset are written by graduates of only twenty schools. When we consider the output of the top-ranked schools against the background of the more than six hundred Ph.D.-granting schools included in the dataset, it makes the relative share of the top schools even more significant.

When ranked by production of affiliated authors, Harvard University is the top school and University of Pennsylvania and University of Chicago round out the top three institutions. We also find significant skew towards a few institutions with the top-twenty institutions accounting for a weighted share of twenty-four percent.

Interestingly, the most prolific authors in any of the five disciplines are not affiliated with any of the top-three schools. We also provide information relating to the contribution of top-quality research by non-school institutions, foreign schools, and the relative share of international and U.S.-educated authors in various business disciplines.

Literature Review

Our literature review examines previous research productivity measures with particular emphasis on the relation to perceptions of school and journal quality and faculty productivity. We acknowledge that any attempt to rank schools or journals and thereby infer quality is potentially hazardous. Any definition of "quality" is subject to potential problems of bias or subjectivity. Hence we endeavor to strike a balance between being sufficiently inclusive, i.e., not ignoring any major specialty areas within each discipline, while also including high-quality articles. The journals included in our dataset have low average acceptance rates[2] and are fairly consistent with very little difference in journal rankings over time (Alexander and Mabry 1994, and Heck and Cooley 2005). We acknowledge that our quality proxy may be imperfect because an article published in a top journal may not necessarily be a top article as Smith (2004) and Schwert (1993) show. However, since there is no accurate way to distinguish among articles in a specific journal, all articles within a journal are given equal weighting.

Perceptions of Journal Quality

For many years, academics and administrators have attempted to rank journals based on some hierarchy of "quality". Zinkhan and Leigh (1999) provide a useful summary of seven indicators of journal quality. They are (1) publisher (e.g., a scholarly society vs. a commercial entity); (2) the reputation of the editor or the members of the editorial board; (3) contributing author reputation; (4) acceptance rate; (5) journal age; (6) journal impact; and (7) knowledge use.

However, despite several years of debates, there is still no universally accepted journal rank, but several journals have earned the distinction of consistently making lists of high-quality journals compiled by various authors. For example, both Borde et al. (1999) and Chung et al. (2001) identify the *Journal of Finance* as the leading finance journal. In economics, Heck (1993) identifies the *American Economic Review* (*AER*) as the top journal. Management scholars, such as Johnson and Podsakoff (1994), point to the *Administrative Science Quarterly* as the top management journal, while Bauerly and Johnson (2005) identify the *Journal of Marketing* as the top marketing journal. Howard and Nikolai (1983) indicate that *The Journal of Accounting Research* and *The Accounting Review* are the top accounting journals.

Chan et al. (2001 and 2004) find that international finance researchers consider U.S.-based journals to be appropriate outlets for their studies. We therefore expect to find significant contributions from foreign authors, which further increase competition for article inclusion and enhance each journal's international appeal.

Tuckman and Leahey (1975) demonstrate that publishing in top-ranked economics journals has significant impact on an author's remuneration. Swidler and Goldreyer (1998) also estimate that the financial impact of a finance publication is significant and depends on professorial rank. Ragan et al. (1999) finds that an article published in *AER* boosts pay by as much as 11 percent, whereas an article in an unranked journal increases pay by at most 1%. Hence, the incentive to publish in top journals goes beyond emotional appeal or self-fulfillment but also carries significant financial rewards. Further, Zivney and Bertin's (1992) finding that "publishing one article per year in *any finance journal* over any prolonged period of time is truly a remarkable feat, met by only 5% of graduates" may be used as justification for rewarding prolific authors.

School Quality

Several periodicals such as *Business Week* and *U.S. News & World Report (USNWR)* offer annual school rankings. However, Dichev (1999) finds that changes in both *Business Week* and *USNWR* graduate business school rankings have a strong tendency to revert. They argue that annual rankings changes are due to aggregations of "noisy" information. In this paper we use contribution of graduates which is less noisy and therefore more reliable.

Empirical research also indicates that graduating from any of the so-called "Ivy-League" schools provide more opportunities both within and outside academia. For example, faculty employment trends indicate that there is a *de facto* pecking order in terms of recruitment. Schools tend to hire graduates from similarly- or higher-ranked universities which in essence determine starting remuneration for newly-minted faculty. Graduates of top schools command higher salaries and there is a significant premium for business disciplines versus languages and literature (Ehrenberg et al. 2006). However, such reports run counter to Long et al. (1998) findings that the status of a graduate's academic origin is less important than one's academic affiliation as a predictor of research productivity. This analysis includes an examination of both Ph.D. school and current affiliation, thereby providing additional evidence as to the relative importance of both Ph.D. school and affiliation in determining research productivity in high-quality journals.

Tauer and Tauer (1984) rank doctoral programs using a method somewhat similar to the method used in this study but with several key differences. Their research rank programs based on contributions to only one journal (*American Journal of Agricultural Economics*) and graduates in one major (agricultural economics). Our research adds significant value by examining more journals (48), more programs (Ph.D. programs in five disciplines), and no limit on graduation year.

Heck (2007) also rank doctoral programs by examining the research productivity of the programs' graduates as well as faculty in four leading finance journals[3] over a fifteen-year span[4]. He includes 91 US finance doctoral programs and concludes that the University of Chicago is the top overall school. Our study is similar in some respects but we do not limit our analysis to US doctoral programs and we use a wider selection of journals. We believe our study therefore provides a more comprehensive ranking of doctoral programs.

Ultimately, school rankings have value as many schools refer to college rankings in their marketing[5] campaigns. Additionally, various periodicals often have tremendous sales success for the issue with their latest college rankings as these issues often are among their best-selling issues. Hence, in recent years there has been a proliferation of college rankings and the attendant never-ending debates.

Faculty Productivity

Although faculty productivity is a much discussed topic, wide disagreement on measurement criteria exists as the factors that determine productivity are difficult to identify with certainty. Hickman and Shrader (2000) investigate the factors that might predict the productivity of new professors using nine different variables[6] but find that the only significant determinant of professor productivity is school ranking. They find that the higher the ranking of the school, the more productive its doctoral graduates and the more productive its faculty members. Buchmueller et al. (1999) finds that research experience in graduate school, graduate school ranking, the graduate school faculty size and publishing frequency, as well as individual demographic characteristics are all indicative of publishing frequency. We try to determine if Hickman and Shrader's (2000) and Buchmueller et al.'s (1999) results still hold true almost two decades later.

Aggarwal et al. (2007) consider research productivity over an extended period by examining only data for authors that publish at least twelve articles in sixty finance journals. Their findings indicate that prolific authors tend to begin publication early in their career and remain very productive after tenure although there is a spike in the years immediately after graduation. Our study provides updated information across five business disciplines regarding the distribution of prolific authorship.

Chan et al. (2005) focus on Asia-Pacific universities by examining articles published in 18 accounting journals from 1991 to 2002. They find significant contributions by universities in that region especially by Hong Kong and Singaporean universities especially in recent years. Interestingly, they also find that the research productivity of the top 20 Asia-Pacific universities compare favorably with their North American counterparts. By including all authors irrespective of their current affiliation or Ph.D.-granting institution, we will be able to determine if Chan et al.'s (2005) finding holds across all five disciplines.

Chan et al. (2009) introduces the concept of a "pedigree" effect in author productivity. Their study includes 21 journals, with a focus on the top three (*Journal of Finance, Journal of Financial Economics*, and *Review of Financial Studies*). As expected, they find that authors graduating from "elite" schools tend to be more productive. If this finding holds in our sample we anticipate significant productivity bias in favor of the top schools across all five disciplines.

Faculty Productivity Measures

Despite the consensus to assess faculty and degree programs regularly, there is little agreement on the method of assessment. However, administrators and researchers usually choose one or a combination of i) journal-quality surveys, ii) counting the number of citations in subsequent articles, or iii) counting the number and/or pages of published articles. Crosta and Packman (2005) propose an additional measure of faculty productivity. They argue that since faculty members have the important responsibility of producing new Ph.Ds. to ensure continued knowledge advancement, success in supervising Ph.D. students should serve as an integral measure of faculty productivity. This argument, although not without some merit, unfairly penalize authors affiliated with non-Ph.D.-granting institutions. Further, in measuring the productivity of a researcher, quality is an essential factor, simply because the productivity of two researchers, one of whom publishes a seminal article in a top-tier journal, while the other publishes in a lower-tier journal with very little impact, is arguably different.

Although survey techniques vary, they all involve similar elements such as either asking potential respondents to rank journals according to some base measure, or to group journals in 'quality' tiers or classes. Hence, although survey results provide a source of estimates of 'quality', Alexander and Mabry (1994) point out that a survey may be biased. Oltheten et al. (2005) attribute differences in researchers' geographic origin, research interests, seniority, and journal affiliation, as predictors of journal-quality perceptions. These factors could possibly taint surveys resulting in biased findings.

Borde et al. (1999) attempts to eliminate survey bias by using a more homogenous group by surveying finance department chairpersons. They still find similar results to Oltheten et al. (2005), i.e., consistency in ranking the top journals and substantial variations among lesser-ranked journals. These findings are not unique to the finance discipline. Howard and Nikolai (1983) while finding that *The Journal of Accounting Research* and *The Accounting Review* consistently rank as the top two accounting journals, they also find little agreement in quality perceptions for journals across specialty areas within accounting. They conclude that perception variances are determined by differences among faculty affiliated to doctoral- and non-doctoral granting institutions as well as faculty rank.

Many administrators determine journal "quality" by the 'acceptance rate', i.e., the percentage of submitted articles published in the journal. The thinking is that low acceptance rates indicate higher screening standards, implying that the published articles are of a higher quality. Such reasoning seems convincing, but other factors such as submission costs, editorial affiliation, country of origin, and circulation may also affect journal submission rates. Further, as administrators increase pressure on faculty to improve research quantity and quality, journals with traditionally low acceptance rates may have the number of submitted articles artificially inflated as poorly written articles, which would not be published in "lesser" journals, will be submitted by authors trying to get a high-quality "hit".

Kacmar and Whitfield (2000) insists that the citation method is more objective than opinion surveys (the paper is either cited or not cited), while still pointing to impact (i.e., quality) but point out that a major limitation is that each citation is awarded equal weight. However, the citation method is not without distracters as Alexander and Mabry (1994) show that counting citations for articles may not be an appropriate quality measure as many articles suffer from self-citation bias. Further, a few prominent researchers and journals dominate citations (Chung et al. 2001 and Cox and Chung, 1991). In addition, citation volume is also driven by article exposure as larger circulation would imply greater visibility and hence a greater likelihood of citation. Additionally, Bonzi (1992) find that citation counts are biased toward older works, since they have had greater exposure over time when compared to more recent works.

The counting method as used by researchers such as Heck and Cooley (2005) and Hasselback et al. (2000) enumerate the number of articles published by authors. Although the counting method appears to be the most objective, Heck and Cooley's (2005) study in which they list the most prolific authors in the finance literature over a 50-year period demonstrates how journal selection impact productivity rankings. Their top-ranked author lists change considerably when the journal list is increased from seven to sixteen and then seven-two journals.

Data

Several hundred journals provide publishing outlets for business-related manuscripts. We draw our inferences by combining several journal lists compiled by previous researchers. Our use of ten journals[7] each for Accounting, Finance, Economics, Management, and Marketing allows us to focus on the top journals in each discipline, while still allowing sufficient breadth to cover major subjects in each discipline. Further, our comprehensive journal list provides authors with more publishing outlets than those afforded in studies reviewed earlier. Also, by considering journals in five business disciplines we can capture any cross-discipline research.

Our list of journals on average have low acceptance rates so the quality of each journal is not only more convenient to establish, but is generally widely accepted. We acknowledge that our quality proxy may be imperfect because an article published in a top journal may not necessarily be a top article as Smith (2004) and Schwert (1993) show. However, most articles in a top journal are of reasonably good quality as evidenced by the fact that the group of top journals is consistent with very little difference in journal rankings in various studies over time (Alexander and Mabry 1994, Borde et al. 1999, and Heck and Cooley 2005).

Following Borokhovich et al. (1995), only articles and notes are included (editorials, comments, and replies are omitted) in developing a list of all authors who publish at least one article in any of the journals between 2000 and 2009.We compile our data from the table of contents for each issue of the journals to identify authors and their affiliated institutions[8]. Table 1 (Panel A) indicates that a total of 25,997 articles written by more than 15,000 authors from over 2,000 unique institutions are included in our dataset. Authors identified with Ph.D. degrees earn their qualifications from approximately 600 schools.

Table 1: Summary Statistics for Articles Published in 48 Elite Business Journals (January 2000 – December 2009)

Panel A: Authors & Institutions

Business Discipline	Total Articles	Total Authors	Total Affiliated Institutions	PhD-Granting Schools	Average Articles per Author	Average Articles per Institution	Average Articles per PhD School
Accounting	2,450	1,387	529	260	1.77	4.63	9.42
Economics	6,138	5,834	1,014	374	1.05	6.05	16.41
Finance	5,603	5,866	1,360	473	0.96	4.12	11.85
Management	6,076	8,080	1,331	526	0.75	4.56	11.55
Marketing	5,730	6,873	1,544	524	0.83	3.71	10.94

We also notice some difference in the number of articles for accounting compared to the other four disciplines. Economics, finance, management, and marketing average 5,887 articles each with all four journals having total articles within 5% of the average. Accounting articles total 2,450 or less than half that for each of the other four disciplines. We account for this discrepancy by calculating a weighted overall rank. U.S. schools account for the majority of articles, but foreign schools and non-school institutions also make significant contributions, especially in economics, finance and marketing (Table 1, Panel B). We use the Internet to obtain data relating to each author's school of Ph.D. and year Ph.D. completed[9].

Table 1: Summary Statistics for Articles Published in 48 Elite Business Journals (January 2000 – December 2009)

Panel B: US versus Foreign Institutions (by author affiliation)

Business Discipline	Articles by US Institutions	% of Contributions	Articles by Foreign Institutions	% of Contributions	Total Articles
Accounting	1,865.5	76.1	584.5	23.9	2,450
Economics	4,266.9	69.5	1,871.1	30.5	6,138
Finance	3,659.7	65.3	1,943.3	34.7	5,603
Management	4,467.3	73.5	1,608.7	26.5	6,076
Marketing	3,949.3	68.9	1,780.7	31.1	5,730

Methodology

We use the counting technique to evaluate research productivity because it is an objective and cost-effective method (Hasselback et al. 2003). We also focus on high-quality articles since we only use the top journals in each discipline. The number of articles that an author publishes in the ten-year period is the *total number of appearances* the author has to his/her credit. To avoid double counting, we follow Heck and Cooley (2005) and calculate an *adjusted number of articles* per author by using weights of 0.5 for two authors, 0.333 for three, 0.25 for four and so on. The number of total and adjusted articles per institution is based on the author's affiliation as indicated in the journal index.[10] We also sort the data using each author's Ph.D.-granting institution to calculate the total and adjusted number of articles per Ph.D.-granting school.

Results

We first present the results from our counting analysis by Ph.D.-granting institution and compare our result with an affiliation-based ranking. The third section of our results present our findings on individual author productivity.

Ph.D.-Granting Schools

The most productive school (based on total adjusted articles weighted for contribution in each discipline) is Massachusetts Institute of Technology (MIT) with Harvard University and University of Chicago second and third respectively (Table 2). MIT's dominance is very consistent as that school places in the top five Ph.D.-granting school in four of the five disciplines with accounting as the exception as MIT falls to 32nd place in that discipline. Further, we find that many schools which dominate in one discipline rank very low in other disciplines. Only four of the overall top-twenty schools appear in the top-twenty for all five discipline-specific lists. This implies that Ph.D. programs concentrate in specific disciplines rather than across all five business disciplines considered. However, there is still a significant concentration of authors as the top-twenty Ph.D.-granting schools account for 44.1% of all adjusted articles.

Although several foreign Ph.D.-granting schools such as University of Oxford and London School of Economics rank highly in one or more disciplines, none make our consolidated top-twenty list of Ph.D.-granting schools.

Affiliation Ranking

Sorting the data by author affiliation indicates that Harvard University is the top overall institution with a weighted share of 2.0% for all five disciplines (Table 3). Interestingly, Harvard University ranks in the top three schools for only economics and finance, but score sufficiently high in the other disciplines to earn the top overall spot. The University of Pennsylvania and University of Chicago rank second and third overall. There is less concentration among the highest-ranked author affiliated schools as the top-twenty institutions account for 24.1% of all articles, significantly less than the 44.1% share when the data is sorted by Ph.D.-granting schools.

As in the case of Ph.D.-granting institutions, US schools dominate the affiliation rankings. Authors affiliated with foreign institutions account for 7,788.3 articles with the largest contribution in finance (3,659.7 or 34.7%). No foreign institution rank in the top-twenty overall, but Hong Kong University of Science & Technology ranks #15 in accounting, London School of Economics ranks #15 in economics, London Business School ranks #13 in finance, and INSEAD ranks #14 in management and #10 in marketing as the top-performing foreign institutions.

Authors affiliated with the Board of Governors of the Federal Reserve System had a weighted contribution that ranked that institution at #19, the only non-school institution ranked in the top twenty. Taken together, all Federal Reserve institutions account for 562.5 adjusted articles (2.2% of all articles) which demonstrate the significant contribution of these institutions to research advancement particularly in economics and finance. Additionally, several other financial institutions rank highly in economics and finance, but no non-school institution rank in the top-twenty in the other three disciplines.

Prolific Authors

The most prolific authors in each discipline substantially out-perform their peers whether we use the number of appearances, pages published, or adjusted article count to measure productivity. For accounting, the top author is Kannan Raghunandan with 29 appearances and 11.1 adjusted articles (Table 4). In economics, John List tops the ranking with 23 appearances and 14.2 adjusted articles (Table 5). Table 6 indicates that the top author in finance is Jeff Madura with 28 appearances and 11.8 adjusted articles while Timothy Judge tops the list in management with 42 appearances and 18.0 adjusted articles (Table 7). Dhruv Grewal is the most prolific author in marketing with 37 appearances and 12 adjusted articles. The performance of the most prolific author in each discipline is truly outstanding when viewed against the fact that in all disciplines, more than 50% of all authors appear only once and the average adjusted article count is less than 1. Further, in all disciplines the vast majority of authors appear only once when we only consider authors holding a Ph.D. degree for the entire ten-year period under review.

Our results also demonstrate that the most prolific authors tend to publish with co-authors as evidenced by the number of appearances versus adjusted article count. Interestingly, the top authors in finance, accounting, management, and marketing are not affiliated with the top overall schools which demonstrate that prolific authorship is determined by more than affiliation or school of Ph.D. We find very few instances of cross-discipline authorship and no single author made any significant impact in more than one discipline.

Further analysis also indicate that many of the top-twenty most prolific authors are associated with foreign institutions (#s 4, 8, and 17 for management, #s 8, 9, 10 and 15 for marketing, #s 2, 4 and 8 for accounting, #s 2, 10, 11, and 17 for finance, and #4 for economics). Considered against the background that many of these authors' primary language is not English, and the additional time and expense of dealing with foreign journals, the accomplishments of these authors are remarkable. Our findings clearly indicate that the most prolific authors are truly exceptional and are driven by factors other than being associated with the best U.S. schools.

Conclusion

We analyze articles published in 10 elite business journals for each of five business disciplines between 2000 and 2009 to investigate the comparative productivity across business disciplines. Our dataset includes 25,997 articles written by over 15,000 authors that are affiliated with more than 2,000 institutions. Our analysis includes data on the Ph.D.-granting schools of authors in the dataset to determine the impact of Ph.D. program on research productivity.

MIT has the most (adjusted) articles at 1,200.6 or a weighted share of 4.6% when the data is sorted by Ph.D.-granting school. We find a significant skew in Ph.D.-school contributions with the top-twenty schools accounting for 44% of all articles. Overall, there is a relatively small number of Ph.D.-granting institutions that are included in the dataset which possibly demonstrates significant differences in program quality. Interestingly, only four of our overall top-twenty schools are also listed among the top-twenty Ph.D.-granting schools in all five disciplines. This indicates the relative degree of specialization among the top schools. We find very few instances of cross-discipline authorship and no single author make a significant impact in more than one discipline.

Our findings provide additional insights that could be useful in determining faculty research targets, as well as contribute empirical evidence to the on-going debate regarding the relative productivity across business disciplines. Further, our findings appear to indicate a continued trend towards co-authorship and very little cross-discipline research.

Notes

1. We use a total of 48 journals because the *Journal of Financial Economics* is included in lists for both finance and economics and *Management Science* appears in our list for management and marketing.
2. The average acceptance rate ranges from a low of 4% for the *Journal of Finance* to a high of 30% for *Journal of Vocational Behavior* according to Cabell's Directories of Publishing Opportunities (2010).
3. The four journals are *Journal of Finance*, *Journal of Financial and Quantitative Analysis*, *Journal of Financial Economics*, and *Review of Financial Studies*.
4. The data covers articles published 1991-2005.
5. For example, Texas A&M May's School of Business' website proudly boasts that their MBA program is 1st U.S. overall (19th in world), their undergraduate business program is 17th public (44th overall) based on various rankings. They also provide several other results from specialized rankings by *The Princeton Review*, the *Wall Street Journal*, and the *U.S. News & World Report* among others. (http://mays.tamu.edu/about-mays/rankings/ Accessed January 28, 2011).
6. Productivity is the dependent variable; with independent variables for rank of doctorate granting school, number of publications listed on the resume, number of presentations made at scholarly meetings, ranking of the school of hire, and dummy variables (with values of 0 or 1) to represent presence of a BA in a technical course like (science, math, etc.), an undergraduate degree in business or economics, a US degree, gender, and if the individual had a Ph.D. when the resume was listed in the resume book.
7. See Appendix for the list of journals covered in this study.
8. Some authors are associated with multiple institutions during the review period. In such cases, we ascribe credit to the author's institution as listed in the respective journal entry.
9. We use *ProQuest*'s digital dissertation abstracts as well as each author's vita and/or profile from their personal or affiliated institution's website.
10. Several authors are included in the count for more than one institution. In our summary tables, we use the last institution in the dataset, which may not be the author's current institution, or the institution to which some articles have been attributed.

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Table 2: Summary Statistics for 5 Disciplines by Ph.D.-Granting Institutions

Overall Rank		Accounting		Economics		Finance		Management		Marketing		Total	
		Articles	**Share**	**Articles**	**Share**	**Articles**	**Share**	**Articles**	**Share**	**Articles**	**Share**	**Articles**	**Share**
1	MIT	21.3	0.9%	560.5	9.1%	255.4	4.6%	173.5	2.9%	190.0	3.3%	1,200.6	4.6%
2	Harvard University	29.7	1.2%	501.9	8.2%	212.8	3.8%	131.9	2.2%	63.0	1.1%	939.2	3.6%
3	University of Chicago	89.7	3.7%	377.9	6.2%	276.2	4.9%	66.7	1.1%	86.3	1.5%	896.9	3.5%
4	Stanford University	78.6	3.2%	313.7	5.1%	154.7	2.8%	250.4	4.1%	227.7	4.0%	1,025.0	3.9%
5	University of Pennsylvania	36.3	1.5%	213.4	3.5%	152.0	2.7%	183.2	3.0%	183.3	3.2%	768.2	3.0%
6	Northwestern University	42.5	1.7%	177.9	2.9%	85.3	1.5%	160.9	2.6%	190.1	3.3%	656.7	2.5%
7	U of California, Berkeley	40.3	1.6%	219.9	3.6%	108.5	1.9%	151.4	2.5%	103.8	1.8%	623.8	2.4%
8	University of Michigan	109.1	4.5%	72.3	1.2%	61.8	1.1%	170.3	2.8%	110.3	1.9%	523.8	2.0%
9	University of Minnesota	37.2	1.5%	205.9	3.4%	38.7	0.7%	157.9	2.6%	74.5	1.3%	514.2	2.0%
10	Princeton University	6.8	0.3%	240.2	3.9%	83.9	1.5%	30.9	0.5%	17.8	0.3%	379.5	1.5%
11	University of Rochester	54.9	2.2%	203.8	3.3%	104.7	1.9%	23.4	0.4%	35.9	0.6%	422.7	1.6%
12	U of Illinois Urbana-Champaign	61.8	2.5%	37.2	0.6%	55.1	1.0%	183.2	3.0%	94.6	1.7%	431.9	1.7%
13	Columbia University	33.5	1.4%	86.9	1.4%	72.0	1.3%	117.9	1.9%	133.2	2.3%	443.5	1.7%
14	Yale University	4.5	0.2%	180.3	2.9%	94.6	1.7%	58.3	1.0%	32.0	0.6%	369.6	1.4%
15	UCLA	13.0	0.5%	94.5	1.5%	137.9	2.5%	82.3	1.4%	77.7	1.4%	405.4	1.6%
16	University of Texas, Austin	72.6	3.0%	25.7	0.4%	68.2	1.2%	80.5	1.3%	128.4	2.2%	375.3	1.4%
17	New York University	27.8	1.1%	66.4	1.1%	142.6	2.5%	78.5	1.3%	74.5	1.3%	389.7	1.5%
18	Cornell University	43.2	1.8%	93.5	1.5%	67.1	1.2%	113.3	1.9%	72.0	1.3%	389.1	1.5%
19	U of Wisconsin, Madison	50.4	2.1%	98.9	1.6%	70.2	1.3%	67.9	1.1%	80.6	1.4%	368.0	1.4%
20	Ohio State University	52.7	2.2%	35.7	0.6%	104.2	1.9%	87.5	1.4%	70.0	1.2%	350.0	1.3%
												11,473.3	**44.1%**

Table 3: Summary Statistics for 5 Disciplines by Author Affiliation

Overall Rank		Accounting		Economics		Finance		Management		Marketing		Total	
		Articles	**Share**	**Articles**	**Share**	**Articles**	**Share**	**Articles**	**Share**	**Articles**	**Share**	**Articles**	**Share**
1	Harvard University	26.3	1.1%	238.7	3.9%	101.9	1.7%	103.7	1.7%	60.1	1.0%	530.6	2.0%
2	University of Pennsylvania	0.0	0.0%	153.7	2.5%	76.7	1.2%	117.8	1.9%	133.0	2.3%	481.1	1.8%
3	University of Chicago	47.8	1.9%	200.4	3.3%	82.5	1.3%	34.2	0.6%	55.1	1.0%	419.8	1.6%
4	New York University	36.5	1.5%	156.9	2.6%	117.1	1.9%	76.1	1.3%	75.0	1.3%	461.5	1.7%
5	Northwestern University	25.8	1.1%	155.9	2.5%	33.9	0.6%	71.9	1.2%	88.8	1.6%	376.2	1.4%
6	Stanford University	36.4	1.5%	139.3	2.3%	45.8	0.7%	90.4	1.5%	67.5	1.2%	379.4	1.4%
7	Columbia University	26.7	1.1%	103.5	1.7%	63.1	1.0%	92.8	1.5%	97.1	1.7%	383.1	1.4%
8	MIT	27.7	1.1%	154.6	2.5%	45.2	0.7%	51.9	0.9%	57.9	1.0%	337.3	1.3%
9	University of Michigan	33.5	1.4%	79.3	1.3%	51.0	0.8%	90.7	1.5%	64.9	1.1%	319.4	1.2%
10	U of California, Berkeley	12.3	0.5%	130.1	2.1%	40.9	0.7%	47.9	0.8%	49.9	0.9%	281.1	1.1%
11	UCLA	13.9	0.6%	108.5	1.8%	65.0	1.1%	49.9	0.8%	57.3	1.0%	294.5	1.1%
12	Duke University	28.9	1.2%	61.7	1.0%	49.8	0.8%	68.7	1.1%	86.3	1.5%	295.3	1.1%
13	University of Maryland	3.9	0.2%	54.3	0.9%	42.1	0.7%	100.1	1.6%	72.3	1.3%	272.7	1.0%
14	Princeton University	1.3	0.1%	126.3	2.1%	29.5	0.5%	20.1	0.3%	2.1	0.0%	179.3	0.7%
15	Pennsylvania State University	27.2	1.1%	43.333	0.7%	39.5	0.6%	92.1	1.5%	50.4	0.9%	252.5	1.0%
16	University of Texas, Austin	38.7	1.6%	43.2	0.7%	49.3	0.8%	63.5	1.0%	60.3	1.1%	255.0	1.0%
17	University of Minnesota	12.4	0.5%	47.783	0.8%	24.0	0.4%	95.1	1.6%	54.4	1.0%	233.7	0.9%
18	U of Illinois, Urbana-Champaign	28.5	1.2%	44.75	0.7%	39.3	0.6%	86.3	1.4%	43.0	0.8%	241.8	0.9%
19	Bd of Governors, Fed Res Sys	0.0	0.0%	81.6	1.3%	84.4	1.4%	0.0	0.0%	0.8	0.0%	166.8	0.6%
20	U of Southern California, LA	34.9	1.4%	41.367	0.7%	28.6	0.5%	45.3	0.7%	69.6	1.2%	219.7	0.8%
												6,380.7	**24.1%**

Table 4: Accounting Journals by Authors (2000 - 2009) [*Last affiliation in dataset which may be different from current affiliation or affiliation when some articles were published.]

Rank	Author Name	Adjusted Articles	Adjusted Pages	Appear-ances	Author Affiliation*	Author PhD School	Grad Year
1	Raghunandan, Kannan	11.1	170	29	Florida International U	University of Iowa	1990
2	Tan, Hun-Tong	8.3	186	18	Nanyang Technological U	University of Michigan	1992
3	Hunton, James E.	7.3	142	19	Bentley College	U of Texas, Arlington	1994
4	Krishnan, Gopal V.	7.2	178	10	City U of Hong Kong	U of North Texas	1986
5	Sansing, Richard C.	7.2	128	13	Dartmouth College	U of Texas, Austin	1990
6	Rama, Dasaratha V.	6.7	104	17	Florida International U	University of Iowa	1990
7	Francis, Jere R.	6.5	161	15	University of Missouri	U of New England	1982
8	Lennox, Clive S.	6.5	144	10	H Kong U of Sc & Tech.	University of Oxford	1998
9	Rajgopal, Shivaram	6.4	234	18	Duke University	University of Iowa	1998
10	Bedard, Jean C.	6.1	144	16	Bentley College	U of Wisconsin, Madison	1985
11	Shevlin, Terry	6.1	163	15	University of Washington	Stanford University	1986
12	Johnstone, Karla M.	5.8	146	13	U of Wisconsin, Madison	U of Connecticut	1997
13	Barth, Mary E.	5.8	143	13	Stanford University	Stanford University	1989
14	Weber, Joseph P.	5.7	152	13	MIT	Pennsylvania State U	2000
15	Wright, Arnold M.	5.7	109	15	Boston College	U of Southern California	1979
16	Arya, Anil	5.5	97	11	Ohio State University	University of Iowa	1991
17	Kadous, Kathryn	5.3	126	11	Emory University	U of Ill, U-Champaign	1996
18	Dhaliwal, Dan S.	5.3	131	16	University of Arizona	University of Arizona	1977
19	Kaplan, Steven E.	5.3	110	12	Arizona State University	U of Ill, U-Champaign	1982
20	Ke, Bin	5.2	148	11	Pennsylvania State U	Michigan State U	1999
	Sub-Total	128.6	2,914	295			
48	Auths with 4 to <5.2 Adj Articles	210.0	5,712	424			
78	Authors with 3 to <4 Adj Articles	259.4	6,786	543			
192	Authors with 2 to <3 Adj Articles	452.6	11,393	939			
602	Authors with 1 to <2 Adj Articles	772.5	19,566	1,592			
1,387	Authors with <1 Adjusted Article	627.0	15,439	1,665			
	Total	**2,450**	**61,810**	**5,458**			

Table 5: Economics Journals by Authors (2000-2009) [*Last affiliation in dataset which may be different from current affiliation or affiliation when some articles were published.]

Rank	Author Name	Adjusted Articles	Adjusted Pages	Appear-ances	Author Affiliation*	Author PhD School	Grad Year
1	List, John A.	14.2	299	23	University of Central Florida	University of Wyoming	1996
2	Acemoglu, Daron	14.1	503	29	MIT	London Sch of Economics	1992
3	Jackson, Matthew O.	9.9	286	22	California Inst of Technology	Stanford University	1988
4	Tirole, Jean M.	9.3	284	16	Ins d'Economie Industrielle	MIT	1981
5	Sandholm, William H.	9.0	218	11	U of Wisconsin, Madison	Northwestern University	1998
6	Shi, Shouyong	8.3	246	11	Indiana University	University of Toronto	1991
7	Andrews, Donald W. K.	8.3	290	11	Yale University	U of California, Berkeley	1982
8	Sargent, Thomas J.	8.1	265	18	New York University	Harvard University	1968
9	Matsuyama, Kiminori	8.0	193	8	Northwestern University	Harvard University	1987
10	Wright, Randall	7.9	161	19	University of Pennsylvania	University of Minnesota	1986
11	Kocherlakota, Narayana R.	7.8	146	11	Stanford University	University of Chicago	1987
12	Samuelson, Larry W.	7.5	196	15	U of Wisconsin, Madison	U of Illinois, U-Champaign	1978
13	Ray, Debraj	7.4	217	16	New York University	Cornell University	1983
14	Newey, Whitney K.	7.2	182	12	MIT	MIT	1983
15	Segal, Ilya R.	7.0	209	10	Stanford University	Harvard University	1995
16	Manski, Charles F.	7.0	141	9	Northwestern University	MIT	1973
17	Horner, Johannes A.	6.9	195	13	Northwestern University	University of Pennsylvania	2000
18	Levine, David K.	6.7	173	14	Washington University	MIT	1981
19	Lagos, Ricardo	6.7	149	11	Fed Res Bk of Minneapolis	University of Pennsylvania	1997
20	Shleifer, Andrei	6.6	214	18	Harvard University	MIT	1986
	Sub-Total	168.0	4,567	297			
51	Auths with 5 to <6.6 Adj Articles	289.2	7,896	485			
97	Authors with 4 to <5 Adj Articles	416.3	11,453	771			
190	Authors with 3 to <4 Adj Articles	631.0	16,989	1,125			
493	Authors with 2 to <3 Adj Articles	1,138.4	29,993	2,042			
1,570	Authors with 1 to <2 Adj Articles	1,923.5	47,644	3,364			
3,413	Authors with <1 Adj Article	1,571.6	38,044	3,952			
	Total	**6,138**	**156,587**	**12,036**			

Table 6: Finance Journals by Authors (2000 - 2009) [*Last affiliation in dataset which may be different from current affiliation or affiliation when some articles were published.]

Rank	Author Name	Adjusted Articles	Adjusted Pages	Appear-ances	Author Affiliation*	Author PhD School
1	Madura, Jeff J.	11.8	240	28	Florida Atlantic University	Florida State University
2	Moshirian, Fariborz	10.6	130	14	University of New South Wales	Monash University
3	Berger, Allen N.	9.7	284	23	Bd of Governors of the Fed Res Sys	UCLA
4	Subrahmanyam, Avanidhar	9.3	255	22	UCLA	UCLA
5	Stulz, René M.	9.3	304	25	Ohio State University	MIT
6	Akhigbe, Aigbe	9.2	171	22	University of Akron	University of Houston
7	Longstaff, Francis A.	8.7	224	17	UCLA	University of Chicago
8	Laeven, Luc	8.6	243	18	IMF	U of Amsterdam
9	Chung, Kee H.	8.5	174	20	SUNY, Buffalo	University of Cincinnati
10	Thornton, Daniel L.	8.0	184	11	Federal Reserve Bank of St. Louis	U of Missouri, Columbia
11	Faff, Robert W.	7.8	148	16	Monash University	Monash University
12	Bali, Turan G.	7.8	226	14	Baruch College, CUNY	City U of New York
13	Titman, Sheridan	7.8	200	18	University of Texas, Austin	Carnegie Mellon U
14	Noe, Thomas H.	7.7	224	15	Tulane University	U of Texas, Austin
15	Schultz, Paul H.	7.5	233	10	University of Notre Dame	University of Chicago
16	Stiroh, Kevin J.	7.5	205	10	Federal Reserve Bank of New York	Harvard University
17	Massa, Massimo	7.4	275	15	INSEAD	Yale University
18	Hasan, Iftekhar	7.4	156	20	Rensselaer Polytechnic Institute	University of Houston
19	Chordia, Tarun	7.3	220	19	Emory University	UCLA
20	Stein, Jeremy C.	6.7	211	15	Harvard University	MIT
	Sub-Total	168.4	4,306	352		
#						
34	Auths with 5 to <6.7 Adj Articles	193.4	5,398.6	408.0		
51	Authors with 4 to <5 Adj Articles	222.2	6,470.1	431.0		
163	Authors with 3 to <4 Adj Articles	548.4	15,499.7	1,084.0		
416	Authors with 2 to <3 Adj Articles	976.2	26,712.1	1,963.0		
1,427	Authors with 1 to <2 Adj Articles	1,780.6	44,773.3	3,444.0		
3,755	Authors with <1 Adj Article	1,713.9	39,715.6	4,384.0		
	Total	**5,603**	**142,875**	**12,066**		

Table 7: Management Journals by Authors (2000-2009) [*Last affiliation in dataset which may be different from current affiliation or affiliation when some articles were published.]

Rank	Author Name	Adjusted Articles	Adjusted Pages	Appear-ances	Author Affiliation*	Author PhD School	Grad Year
1	Judge, Timothy A.	18.0	281	42	University of Florida	U of Illinois, U-Champaign	1990
2	Luo, Yadong	13.4	267	17	University of Miami	Temple University	1995
3	Shane, Scott A.	9.5	150	19	Case Western Reserve U	University of Pennsylvania	1992
4	Lievens, Filip	8.7	104	21	Ghent University	Ghent University	1999
5	Allen, Tammy D.	8.6	136	19	U of South Florida	University of Tennessee	1996
6	Colquitt, Jason A.	8.4	129	20	University of Florida	Michigan State University	1999
7	Westphal, James D.	8.3	230	18	U of Michigan, Ann Arbor	Northwestern University	1996
8	Greve, Henrich R.	8.1	171	13	INSEAD	Stanford University	1994
9	Hitt, Michael A.	7.4	145	25	Arizona State University	U of Colorado, Boulder	1974
10	Sackett, Paul R.	7.0	73	19	University of Minnesota	Ohio State University	1979
11	Morgeson, Frederick P.	7.0	96	19	Michigan State U	Purdue University	1998
12	Cannella Jr., Albert A.	6.9	120	18	Arizona State University	Columbia University	1991
13	Hambrick, Donald C.	6.8	132	15	Columbia University	Pennsylvania State U	1979
14	Tracey, Terence J.G.	6.7	150	12	Arizona State University	U of Maryland, College Park	1981
15	Ilies, Remus	6.7	89	18	Michigan State U	University of Florida	2003
16	LePine, Jeffery A.	6.5	104	15	University of Florida	Michigan State University	1998
17	De Dreu, Carsten K.W.	6.5	106	12	University of Amsterdam	University of Groningen	1993
18	Carpenter, Mason A.	6.5	113	13	U of Wisconsin, Madison	University of Texas, Austin	1997
19	Harrison, David A.	6.2	121	17	Pennsylvania State U	U of Illinois, U-Champaign	1988
20	Flynn, Francis J.	6.2	94	10	Columbia University	U of California, Berkeley	2001
	Sub-Total	163.4	2,810	362			
29	Authors with 5 to <6.2 Adj Arts.	159.1	2,745	363			
57	Authors with 4 to <5 Adj Articles	249.8	4,368	510			
120	Authors with 3 to <4 Adj Articles	406.6	6,827	903			
370	Authors with 2 to <3 Adj Articles	861.4	15,216	1,863			
1,403	Authors with 1 to <2 Adj Articles	1,744.2	30,599	3,624			
6,081	Authors with <1 Adj Article	2,491.5	42,261	7,242			
	Total	**6,076**	**104,827**	**14,867**			

Table 8: Marketing Journals by Authors (2000-2009) [*Last affiliation in dataset which may be different from current affiliation or affiliation when some articles were published.]

Rank	Author Name	Adjusted Articles	Adjusted Pages	Appear-ances	Author Affiliation*	Author PhD School	Grad Year
1	Grewal, Dhruv	13.0	147	37	Babson College	Virginia Polytechnic Inst	1989
2	Tellis, Gerard J.	12.4	185	27	U of Southern California	U of Michigan, Ann Arbor	1993
3	Chernev, Alexander	12.3	138	14	Northwestern University	Duke University	1997
4	Homburg, Christian	11.5	193	29	University of Mannheim	University of Karlsruhe	1988
5	Chintagunta, Pradeep K.	11.5	177	28	University of Chicago	Northwestern University	1990
6	Kumar, V.	10.9	172	28	Georgia State University	U of Texas, Austin	1985
7	Luo, Xueming	10.3	137	19	SUNY, Fredonia	Louisiana Tech U	2003
8	Wedel, Michel	9.2	122	23	University of Groningen	U of Wageningen	1990
9	Chang, Chingching	9.0	115	9	National Chengchi U	U of Wisconsin, Madison	1996
10	Cowley, Elizabeth J.	8.8	78	11	U of New South Wales	University of Toronto	1997
11	Zinkhan, George M.	8.6	98	22	University of Georgia	University of Michigan	1981
12	Janiszewski, Chris	8.4	112	20	University of Florida	Northwestern University	1987
13	Simonson, Itamar	8.1	119	17	Stanford University	Duke University	1987
14	Pauwels, Koen H.	7.7	125	18	Dartmouth College	UCLA	2001
15	Verhoef, Peter C.	7.5	124	18	Erasmus University	Erasmus University	2001
16	Ariely, Dan	7.5	94	17	Duke University	Duke University	1998
17	Krishna, Aradhna	7.3	113	16	U of Michigan, Ann Arbor	New York University	1989
18	Jap, Sandy D.	7.2	110	11	Emory University	University of Florida	1995
19	Dhar, Ravi	7.0	78	18	Yale University	U of California, Berkeley	1992
20	Donthu, Naveen	7.0	85	17	Georgia State University	U of Texas, Austin	1986
20	Inman, J. Jeffrey	7.0	90	17	University of Pittsburgh	U of Texas, Austin	1990
	Sub-Total	192.2	2,610.1	416.0			
18	Authors with 6 to <7 Adj Articles	115.1	1,629	248			
27	Authors with 5 to <6 Adj Articles	145.8	2,011	330			
68	Authors with 4 to <5 Adj Articles	296.7	4,064	692			
138	Authors with 3 to <4 Adj Articles	465.4	5,992	1,000			
354	Authors with 2 to <3 Adj Articles	824.9	10,777	1,809			
1,252	Authors with 1 to <2 Adj Articles	1,538.6	19,078	3,157			
4,995	Authors with <1 Adj Article	2,151.4	26,971	5,889			
	Total	**5,730**	**73,130**	**13,541**			

Appendix: Journal list. [*Acceptance rate taken from Cabell's Directory of Publishing Opportunities 2010 (Accessed January 2011).]

Journals	Acceptance Rate*
Accounting	
Accounting Horizons	15 – 16%
Accounting, Organizations & Society	12 – 15%
Accounting Review	11 – 15%
Auditing: A Journal of Practice & Theory	11 – 20%
Behavioral Research in Accounting	20%
Contemporary Accounting Research	11%
Journal of Accounting, Auditing & Finance	10%
Journal of Accounting & Economics	11 – 20%
Journal of Accounting Research	14%
Journal of the American Taxation Association	11 – 20%
Economics	
American Economic Review	6 – 10%
Econometrica	10%
International Economic Review	11 -20%
Journal of Economic Theory	6 -10%
Journal of Financial Economics	n/a
Journal of Monetary Economics	6 – 10%
Journal of Political Economy	9%
Quarterly Journal of Economics	6 – 10%
Review of Economic Studies	5%
Review of Economics & Statistics	15%
Finance	
Financial Management	11 – 20%
Financial Review	6 – 10%
Journal of Banking and Finance	6 – 10%
Journal of Finance	4%
Journal of Financial & Quantitative Analysis	9%
Finance (continued)	
Journal of Financial Economics	n/a
Journal of Financial Research	9%
Journal of Financial Services Research	7 – 12%
Journal of Money, Credit, & Banking	15%
Review of Financial Studies	8%
Management	
Academy of Management Journal	6%
Academy of Management Review	n/a
Administrative Sciences Quarterly	9 – 11%
American Sociological Review	10%
Journal of Applied Psychology	10%
Journal of Management	12%
Management Science	7%
Organizational Behavior & Human Decision Processes	10 – 15%
Journal of Vocational Behavior	21 – 30%
Strategic Management Journal	n/a
Marketing	
Journal of Advertising	11 – 20%
Journal of Advertising Research	25%
Journal of Business Research	6 – 10%
Journal of Consumer Research	10%
Journal of Marketing	11%
Journal of Marketing Research	12%
Journal of Retailing	15%
Journal of the Academy of Marketing Science	6%
Management Science	7%
Marketing Science	13 – 14%

Low P/E Investing – A Tribute to John Neff

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Abstract

In a tribute to investment professional John Neff, we reassess the merits of low P/E investing relative to other strategies and investigate whether these risk-return characteristics have changed through time. We document an incredibly linear relation between P/E ranks and returns, as well as a strong seasonality pattern in the low P/E strategy. Overall, our results show low P/E investing as a superior investment style before, during and after John Neff's tenure. As the first fund manager to press its advantage for the benefit of his shareholders, Neff rightfully deserves his place among the great investors of the 20th century.

Introduction

The low P/E experience of John Neff, America's greatest low P/E investor, was slightly different during his tenure than it is for current low P/E investors. Neff is considered to be one of the top investors of the twentieth century, with a record that continues to be extremely hard to beat. He cemented his reputation as the head of Vanguard's Windsor fund (originally a Wellington Management fund), by turning it into the best performing and largest mutual fund of its time – during Neff's 31 years (1964 to 1995), Windsor beat the S&P500 by an average of 3.15 percentage points each year. Charles Ellis refers to him as the "investments profession's investment professional" (Neff & Mintz, 1999) and plainly states that "nobody has ever managed a large mutual fund so very well for so long. And no one is likely to ever do so again."

John Neff received many labels during his storied career. Some think of him as a value investor, while others see him as a contrarian, but he actually prefers the label of 'low P/E investor' (Neff and Mintz, 1999). To be exact, his investment style would really fall more into the category of "low P/E investing plus", since he considered several other elements along with low P/E. As described in his book, the "pluses" that his style incorporated relative to pure low P/E investing include fundamental growth in excess of 7 percent; yield protection/enhancement; superior relationship of total return to P/E paid; no cyclical exposure without compensating P/E multiple; solid companies, in growing fields; and an overall strong fundamental case. Some of these factors are a little subjective and probably help some investment historian classify John as a value investor, but he was very clear on the foundation of his philosophy.

While in no way can we attempt to replicate John Neff's strategy, this paper analyzes the basics of a more generic, pure low P/E investment strategy. Today the price earnings ratio, along with other valuation tools, is frequently used by financial analysts to evaluate stocks. The conventional wisdom is that there is a strong relationship between a firm's price-earnings ratio and its fundamentals, such as earnings prospect, risk, and dividend policy. Although the seeds of this investment style were planted all the way back in 1940 by Graham and Dodd, it is very interesting to note that this conventional wisdom was by no means accepted when John Neff started working his magic in 1964 – decades before value and growth became popular investment styles. In reassessing the merits of the low P/E investment strategy today, as opposed to Neff's period, we hope to bring the reader some appreciation of the essence of an investment style that was years ahead of its time.

In summary, our results re-emphasize the strength of low P/E investing. We find evidence of an "incredible linearity" between P/E rank and returns, supporting Neff's suggestion of using P/E as an investment yardstick. We also document a strong seasonality pattern in the low P/E strategy, where most of the returns are realized at the turn of the year. This pattern seems to be more pronounced during Neff's tenure than outside his tenure. Compared with alternative investment strategies (high P/E, small/large capitalization, high dividend etc), low P/E investment maintains top ranking across the different sample periods incorporated in our analysis.

The general organization of the paper proceeds as follows. Section I provides a brief review of the history of Low P/E investing in the literature. Section II describes the data and methodology. Section III presents our results by documenting some of the fundamental characteristics of the low P/E strategy during John Neff's tenure as opposed to other time periods. Section IV concludes.

Low P/E Strategies in the Academic Literature

Trading strategies related to earnings have a long tradition in the investment community and have stirred a lot of interest in academic research. As mentioned in the introduction, John Neff was one of the first to follow some form of earnings based strategy and successfully apply it, way before it became the overly studied topic that it is today. However, advice to buy stocks that sell at low earnings multiples can be traced back to Graham and Dodd (1940). Following that, a wealth of academic literature argued that earnings-related variables (such as P/E ratios) are proxies for expected returns (see Nicholson (1960), Ball (1978), Basu (1977) for some early representative examples). Since the findings in all these early studies—basically saying that low P/E stocks provide higher risk adjusted returns—are inconsistent with the hypotheses of CAPM and efficient markets, they opened the door to a long line of research into the potential explanations for the observed empirical phenomenon. This issue has not yet been settled. Interestingly though, while the academic establishment was just barely beginning to document and study the reasons behind the P/E effect, John Neff had been successfully applying these principles at the Windsor fund since 1964!

Academic interest brought to light numerous empirical facts related to P/E and its connection with expected returns. Given that P/E is just one of the valuation ratios computed using price per share (others are size, CF/P, P/B, etc.), the main question raised was whether these ratios capture independent effects or whether results document a single price common denominator effect. In particular, several studies have focused on size and earnings/price ratio and whether one effect subsumes the other, with contradictory results. For example, Reinganum (1981) argues the size effect is subsumed by the E/P effect, while Basu (1983) claims the exact opposite. Cook and Rozeff (1984) employ a comprehensive battery of tests and argue that both effects are at work. While it is possible that market value and earnings/price ratios measure separate aspects of a single underlying effect, to this date the literature has not agreed on the source of such commonality. Empirical data suggest that, while there is a high degree of commonality among the effects, the price based valuation ratios are not all proxies for the same underlying characteristic (see Hawawini and Keim (1998) for a good review of these studies).

The other interesting avenue of research deals with the seasonality of the P/E effect. As Hawawini and Keim (1998) summarize the literature, they point out that a potential source of the significant correlation of the premia of all price ratios can be that these effects are most pronounced in January. Specifically, the average premia during January tend to be positive and are usually significantly larger than the average premia measured during the rest of the year. This seasonality was documented specifically for the E/P premia by Cook and Rozeff (1984) and by Jaffe et al (1989).

Obviously, the general focus in the academic literature is to find the explanation behind the low P/E effect. In general, academic discussion of low P/E investing is incorporated under broad topics such as "value investing" or "contrarian strategies." While neither one of these completely encompasses the nuances of low P/E investing, it is interesting to comment on the general development of the meaning and implications of "value strategies" and "contrarian strategies" in the academic literature.

In their seminal paper, Fama and French (1992) test several price based variables and their explanatory power on returns and conclude that adding book to market, essentially the inverse of the P/B ratio, to the Fama-MacBeth regressions kills the explanatory power of the E/P variable, among others, for stock returns. Starting with their paper and the use of the now ubiquitous Fama French three-factor model, the academic community generally thinks of book-to-market rather than price-to-earnings ratios in identifying value versus growth strategies. Several explanations have been proposed for the value premium in the literature – notably, Davis et al (2000) identify four common stories: 1) the value premium is sample specific; 2) the value premium is a reflection of compensation for risk, albeit hard to identify specifically, in a multifactor version of Merton's ICAPM; 3) it is due to investor overreaction to firm performance (behavioral story); and 4) it traces back to the value characteristic, not risk (also behavioral story). Although our purpose here is by no means to disentangle the merits of the various explanations proposed in the literature, these alternative explanations are not necessarily mutually exclusive and either one of them may apply to the low P/E effect.

Another thing that it is interesting to comment on is the inclusion of the P/E effect in the broad "contrarian investing" category. Most of the early researchers discussed a low-P/E hypothesis, which was summarized by Basu as follows: "Investors' original expectations of future growth of revenues and earnings for both high and low P/E stocks are overstated, leading to exaggerated optimism for the high-P/E and over pessimism for the low-PE group". Later, Dreman and Berry (1995) proposed the investor-overreaction hypothesis, a broader version of the low-P/E hypothesis, to explain the success of other contrarian strategies, in addition to low P/E.

Although there is definitely huge overlap between the notions of "value," "contrarian" and "low P/E" investing, it is important to recognize that there are differences that distinguish low P/E as an investment strategy. In his book, John Neff is very astute about these nuances and clearly makes the point when emphasizing that he would not categorize himself either as a "contrarian" or as a "value" investor, but rather a "low P/E investor." For this reason, we focus in particular on low P/E in our analysis, versus the book-to-market generally used for the value strategies in the academic literature.

Data and Methodology

We draw our data from two main sources. First, we use Kenneth French's website as a source for portfolio returns for various investment styles (i.e. portfolios of stocks sorted based on various characteristics such as P/E, P/CF, MV etc). This data consists of all NYSE, AMEX, and Nasdaq common shares of firms that are included in both CRSP and COMPUSTAT. ADRs and Closed End Funds are excluded from the analysis. We also obtain the three Fama French risk factors from the same website (MKTRF, HML, SMB). Our analysis is maintained at monthly level and covers the period from July 1951 to December 2009. Second, since our focus is on Neff's extraordinary performance, we use the monthly returns of the Windsor fund (VWNDX) during his tenure (July 1964 to October 1995) as a benchmark. Most of our analysis centers around a comparison between the risk/return profile of the Windsor fund during the Neff period and a general Low P/E (or other investment styles) strategy outside of (or including) this period.

In an attempt to briefly summarize our data we first provide descriptive statistics on the P/E portfolios provided on French's website.[1] Portfolios are formed at the end of each June using NYSE breakpoints, and contain all NYSE, AMEX, and NASDAQ stocks for which we have ME for December of t-1 and June of t, and earnings before extraordinary items for calendar year t-1. For ease of analysis, we focus on the monthly returns of quintile portfolios (firms with negative earnings are excluded from the ranking of the five groups). In Table 1 we present the classic performance statistics for the P/E quintile portfolios – Panel A includes the whole period of our analysis, while panel B presents the same numbers for the Neff period (additional to the P/E quintiles, panel B also includes the standard performance measures applicable to the Windsor fund during John Neff's tenure).

We provide a brief description of the performance measures used in Table 1. The monthly average and standard deviations are the classic statistical measures calculated over monthly raw returns. Beta represents covariance with the market, and is ubiquitous in the literature as a measure of systematic risk. The Sharpe ratio is a common reward-to-variability measure and it is calculated as excess return per unit of total risk (standard deviation) – this ratio is used to characterize how well each particular asset compensates the investor for the risk taken (the higher the number the better). In conjunction to the Sharpe ratio, Jensen's Alpha is also widely used to evaluate mutual fund and portfolio managers' performance. This measure captures the "abnormal return" in excess of the risk adjusted return predicted by the CAPM model – it is basically calculated as the difference between the realized return and the expected return using CAPM. Positive alpha investment opportunities are highly sought.

In Panel A of Table 1, we can see that on a pure return basis, the low P/E portfolio has the highest returns, and the high P/E portfolio has the lowest returns. Somewhat satisfying is the relative systemic nature and continuity of the pattern. Returns decline by approximately 0.13% as one moves from the middle P/E group (Q3) to the middle-high (Q4) and to the high P/E group (Q5). The differences in returns going in the other direction (Q3 to Q2 to Q1) are slightly higher, in the neighborhood of 0.17%. However, this slight asymmetry disappears when computing average percentage difference as one moves up the quintile – the percentage differences between the quintiles, as one moves from the high P/E group to the low P/E group are 15.8%, 13.75%, 14.47% and 14.55%, respectively. Interestingly, this pattern is not repeated in the risk measures (standard deviation and beta).

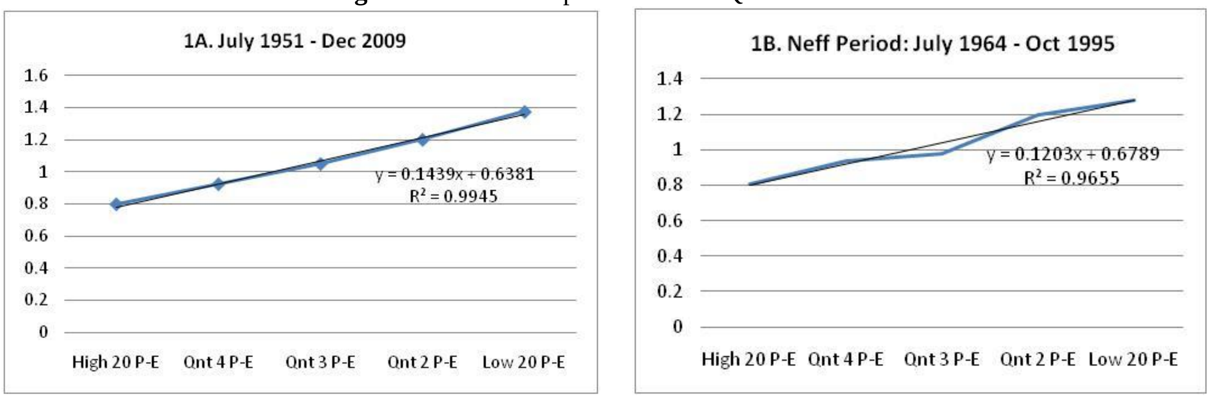
Table 1: Performance Statistics for Returns on Price-Earnings Quintiles Portfolios

Panel A: Whole Period (July 1951 – December 2009)						
	Q5 (High P/E)	**Q4**	**Q3**	**Q2**	**Q1 (Low P/E)**	
Monthly Average	0.797	0.923	1.050	1.202	1.377	
Standard Deviation	4.942	4.188	4.251	4.245	4.816	
Beta	1.089	0.912	0.915	0.893	0.986	
Sharpe	0.081	0.126	0.154	0.190	0.204	
Jensen's Alpha	-0.189	0.033	0.158	0.323	0.447	
Panel B: Neff Period (July 1964 through October 1995)						
	Q5 (High P/E)	**Q4**	**Q3**	**Q2**	**Q1 (Low P/E)**	**Windsor Fund**
Monthly Average	0.805	0.939	0.981	1.196	1.278	1.178
Standard Deviation	5.089	4.514	4.411	4.325	4.837	4.470
Beta	1.098	0.991	0.964	0.925	0.991	0.922
Sharpe	0.053	0.090	0.101	0.153	0.154	0.144
Jensen's Alpha	-0.171	0.006	0.058	0.290	0.345	0.272

Comparing the Neff period (Panel B) with the overall period (Panel A), we see that the Low P-E group still had the highest average monthly return of 1.27%. It also has the best Sharpe ratio and Jensen's Alpha. Looking at the numbers for the Windsor fund, we can identify the "pluses" in Neff's "low P/E plus" strategy. Although the Windsor Fund's average monthly return is not quite as high as a pure low P/E strategy (1.178% vs 1.278%), the standard deviation of his returns is also relatively lower.

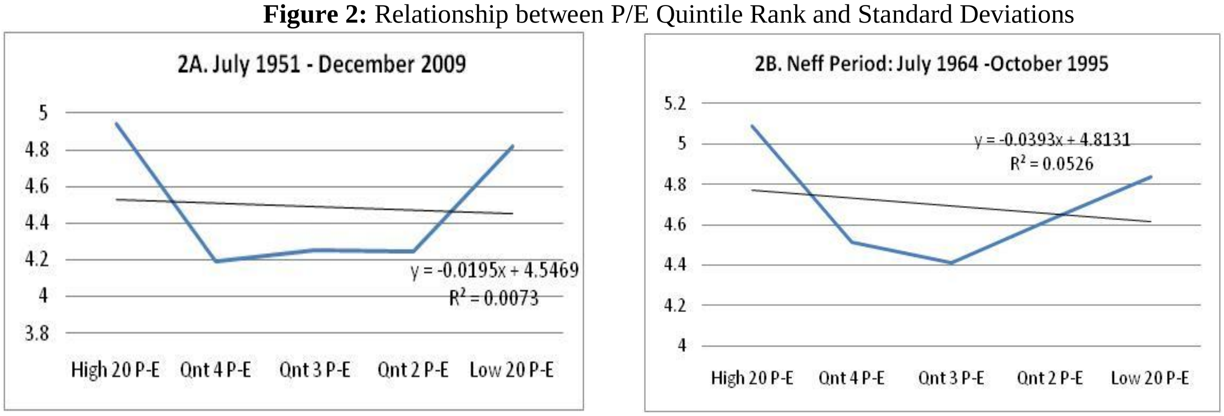
To get a better idea of the P/E relation with return and risk, we provide a graphic description of the numbers in Table 1. We graph the relation between P/E rankings and returns and standard deviations in Figure 1 and Figure 2, respectively, for both the whole period and the Neff Period. Interestingly, the relationship between quintile rankings and returns looks eerily linear over the entire period (Figure 1A). Indeed, the result is so linear that we will dub the phenomena, "incredible linearity." This linearity is consistent with the P/E being an important pricing factor. Arbitrage pricing theory suggests that if a factor is priced that factor will be priced linearly (Ross 1976). For Neff's period, the linearity is slightly less pronounced but the R squared of the regression is still incredibly high.

Figure 1: Relationship between P/E Quintile Rank and Returns



The relationship between standard deviation and P/E quintile rank is nowhere near that linear, neither for the whole period, nor for Neff period in particular.

Figure 2: Relationship between P/E Quintile Rank and Standard Deviations



Standard deviation follows a U-shaped pattern (from Table 1 we can see that this is actually the common pattern for all the other presented measures of risk). This is consistent with the investigation of risk patterns in the B/M studies in the literature, which also find a U-shaped pattern with regard to risk. Some investigators have claim that this implies that the value related phenomena is not related to risk. Overall these figures point more towards a Fama and French (2007) convergence explanation of value phenomena, versus a completely risk based explanation (the systematic pattern in returns is not mirrored by a systematic patter in common risk measures). Neff liked to point out that a low P/E strategy was not necessarily a high risk strategy.

Overall, the descriptive statistics presented above point towards the low P/E strategy ranking very high in terms of its risk-return profile, both during and outside Neff's tenure as a fund manager. We continue investigating its characteristics in the following section.

Empirical Analysis of Low P/E Investing

One of the points John Neff likes to make in his book pertains to the fact that that low P/E investing is a unique form of value investing (Neff and Mintz 1999). He also suggests that the P/E ratio is appropriately the most important yardstick that an investor can use to evaluate stocks. Following on this proposition, we jump start our analysis by comparing a number of different investment strategies in terms of their risk-return profiles. We compare low P/E with a growth strategy (high P/E), a small size strategy, a large capitalization, as well as a high dividend strategy (we obtain returns for all of these investment styles portfolios from French's website[2]). It is important to note that for purposes of comparison we allow only one dimension in each strategy. During the tenure of Mr. Neff, investment funds went by the name of "growth" and "growth and income" etc. Our illustration is intended to show why the Windsor fund with its low P/E philosophy was the dominant fund of its time period.

Table 2: Comparison of Performance Statistics for Various Investment Styles

Panel A: Whole Period (July 1951 – December 2009)					
	Value (Low P-E)	**Growth (High P-E)**	**Small Size**	**Large Size**	**High Dividend**
Monthly Average	1.377	0.797	1.155	0.907	1.041
Standard Deviation	4.816	4.942	5.974	4.168	4.097
Beta	0.986	1.089	1.126	0.950	0.734
Sharpe	0.204	0.081	0.127	0.123	0.158
Jensen's Alpha	0.447	-0.189	0.149	-0.003	0.247
Panel B: Neff Period (July 1964 through October 1995)					
	Value (Low P-E)	**Growth (High P-E)**	**Small Size**	**Large Size**	**High Dividend**
Monthly Average	1.278	0.805	1.190	0.991	1.048
Standard Deviation	4.837	5.089	6.224	4.000	3.815
Beta	0.991	1.098	1.187	0.943	0.720
Sharpe	0.154	0.053	0.105	0.114	0.135
Jensen's Alpha	0.345	-0.171	0.177	-0.019	0.224

As in Table 1, we include two panels: Panel A describes the standard performance statistics for various investment strategies during the whole period, while Panel B repeats this analysis during the period John Neff was an active money manager. We include the classic measures of risk (standard deviation, beta) and performance (Sharpe ratio, Jensen's Alpha) described in Table 1. Ranking the strategies in terms of return, the low P/E had the highest average return over the entire period, followed by the small size, and then the high dividend. The high P/E, sometimes called glamour or growth, performed the worst. In terms of risk, the lowest risk strategy was the high dividend strategy, while the small size was the highest risk strategy. The low P/E strategy had the highest Jensen's Alpha (i.e. the highest abnormal return relative to CAPM risk adjusted expected return).

The next question is the relative performance of the P/E strategy over time. To examine that further we calculate Jensen's alphas for various sample periods and present the results in Table 3 (since Jensen's alphas are calculated relative to CAPM, we also include the beta of each sample period as a benchmark for the risk adjusted return). We can see that the Jensen's alpha for the low P/E strategies was actually better in the periods after John's tenure. Not only was low P/E investing a very good strategy during John's tenure but it also tends to be a very strong performer both before and after John's tenure at Windsor.

Table 3: Jensen's Alpha for Low P-E Strategies over Different Time Periods

Time Period	**Jensen's Alpha**	**Beta**	**Adj. R2**	**F-Value**
July 1951 – Dec 2009	0.447	0.986	0.781	2507
July 1964 – Oct 1995	0.345	0.991	0.811	1611
Nov 1995 – Dec 2009	0.460	0.914	0.715	427
July 1951 – June 1964	0.583	1.142	0.828	633

Although widely used, the CAPM has been highly criticized for its inability to explain the size and book-to-market patterns in returns. Hence, a more robust way to investigate abnormal returns is to look at risk adjusted returns after considering the Fama French three-factor model. If it is the case that low P/E is just a manifestation of size or book-to-market, we should see zero (statistical) alphas coming from a Fama French model. Over different time periods we investigate whether the low P/E strategy might be described by combining a low market to book and a small size portfolio. We run the model over several sample periods and we present the results in Table 4 (panel A). Given that John Neff's strategy is not necessarily a pure low P/E strategy, we also run the same regression using the return of the Windsor funds during his tenure and present the results in Panel B. Numbers presented in parentheses are t-statistics.

Table 4: Risk Adjusted Returns based on Fama French Model

Time Period	FF Alpha	MKT-RF	SMB	HML	Adj. R2
Panel A: Low P/E portfolio returns as dependent variable					
July 1951 – Dec 2009	0.14	1.06	0.17	0.62	0.88
	(2.23)	*(68.22)*	*(7.48)*	*(23.71)*	
July 1964 – Oct 1995	-0.01	1.07	0.23	0.62	0.92
	(-0.13)	*(58.78)*	*(8.81)*	*(21.08)*	
Pre Neff	0.84	1.10	0.05	0.47	0.84
	(2.99)	*(27.99)*	*(0.59)*	*(6.10)*	
Post Neff	0.17	1.00	0.12	0.53	0.83
	(1.00)	*(28.06)*	*(2.52)*	*(10.68)*	
Panel B: Windsor Fund returns during John Neff's period as dependent variable					
July 1964 – Oct 1995	0.063	0.969	0.127	0.368	0.87
	(0.73)	*(44.52)*	*(4.03)*	*(10.50)*	

Over the full sample period the low P/E is not fully explained by the traditional Fama French factors. However, during Neff's tenure a low P/E strategy is mostly explained by combining a low market to book strategy with a small size strategy. An early literature asked the question whether the value effect and size effect were one anomaly or two. Our research helps answer this question by indicating that the answer might be period specific. During Neff's tenure a low P/E strategy is mostly explained by combining a low market to book strategy with a small size strategy. It seems to be that the P/E effect includes some size effect during the Neff tenure. Outside the Neff tenure the size contribution to the P/E effect are much weaker. Looking at Panel B of Table 4, we can see that the Windsor Fund returns are again not identical with the low P/E portfolio – although not statistically significant, Windsor Fund provides a positive excess return (alpha) relative to the Fama-French three factors. It also has lower loadings on SMB and HML than the pure low P/E portfolio. It is important to note that we are comparing actual firms' returns versus theoretical portfolios with no overhead or trading fees. In reality, it is very difficult for practitioners to capture the value premium. For example, Loughran points out "there is a remarkable conflict between finance literature and actual money market performance." He says "the finance literature declares that value firms have reliably higher realized returns than growth firms. However, the realized returns on value and growth money managers are not materially different" (Loughran 1997). As evidence, Loughran cites Malkiel (1995) who reports "in a sample of mutual funds without survivorship bias, that growth funds have an average annual return of 15.81%, during 1982-1991, while growth and income (i.e., value) funds have an average return of 15.97%."

A second important issue that we look at is the seasonality of the low P/E strategy in general, and of the Windsor funds returns in particular. As mentioned in the literature review, low P/E strategies have been documented to exhibit stronger returns in January, and we start by investigating whether that is the case with our data.

We calculate average raw returns over the five P/E quintiles portfolios for each month of the year for various time periods. We identify the month of January versus the average monthly returns during the year and include the results in Table 5. The last line of each panel presents the difference between January and the corresponding monthly returns for all the months in the year (i.e. testing whether January is statistically different from the rest of the year). The monthly returns considered for each year include the particular January month. Numbers in parentheses are t-statistics testing whether the difference is statistically different from 0. Panel A includes the whole period under study; panels B and C isolate the Neff and non-Neff periods, respectively. For comparison purposes, we also calculate the monthly returns and the January returns for the Windsor fund during John Neff's tenure and include these in the last column of panel B.

The results show that during Neff's tenure low P/E investing was a highly seasonal affair. January returns are very strong during John's tenure. Apparently, the good companies received "the attention they deserved" in January. For the period when John Neff managed the Windsor Fund, 1964-1995, the Low P/E group outperformed the rest, and the month of January overwhelmingly provided the highest returns to the fund, accounting for 4.7%. Coming in second, at less than half of the return, was the month of December, which provided 1.9% (results for each individual month are not tabulated, but are available upon request). The total yearly average return for the Low P/E group was 15.42%, compared with 11.82% for the High P/E group and 9.72% for the Middle 20 P/E group.

Table 5: Monthly Performance of P/E Quintile Portfolios

Month	Q1 (Low PE)	Q2	Q3	Q4	Q5 (High PE)	Windsor Fund
Panel A: Whole Period (July 1951 - December 2009)						
January	2.882	1.968	1.328	1.141	0.536	
	(1.89)	*(1.19)*	*(0.50)*	*(0.41)*	*(-0.41)*	
Monthly Avg	1.365	1.190	1.029	0.904	0.794	
	(6.02)	*(6.13)*	*(5.82)*	*(5.25)*	*(3.88)*	
Jan – Monthly Avg	1.517	0.777	0.300	0.237	-0.258	
	(1.89)	*(1.19)*	*(0.50)*	*(0.41)*	*(-0.41)*	
Panel B: Neff Period (July 1964 - October 1995)						
January	4.707	3.355	2.215	1.700	1.376	3.556
	(2.85)	*(2.21)*	*(1.28)*	*(0.80)*	*(0.56)*	*(2.14)*
Monthly Avg	1.298	1.198	0.987	0.947	0.819	1.191
	(4.33)	*(4.89)*	*(4.16)*	*(4.21)*	*(3.18)*	*(5.01)*
Jan – Monthly Avg	3.410	2.157	1.228	0.752	0.557	2.365
	(2.85)	*(2.21)*	*(1.28)*	*(0.80)*	*(0.56)*	*(2.49)*
Panel C: Non-Neff Period (July 1951 - June 1964, November 1995 - December 2009)						
January	0.786	0.374	0.310	0.500	-0.430	
	(-0.74)	*(-1.08)*	*(-1.18)*	*(-0.56)*	*(-1.63)*	
Monthly Avg	1.443	1.182	1.077	0.855	0.765	
	(4.12)	*(3.78)*	*(3.99)*	*(3.17)*	*(2.32)*	
Jan – Monthly Avg	-0.657	-0.807	-0.767	-0.355	-1.195	
	(-0.74)	*(-1.08)*	*(-1.18)*	*(-0.56)*	*(-1.63)*	

Surprisingly, this seasonal phenomenon is not as prominent in the non-Neff period. Does this suggest that portfolio rebalancing strategies have changed? In any event, seasonality in the P/E does appear to be period specific. The magnitude of seasonality during the Neff period is shockingly high indicating that this is not just an aberration that has been "found" by sample selection. The t-statistic in Panel B of Table 5 comparing the January versus the Yearly Monthly during the Neff period is 2.85, which is extremely significant. Indeed, during the non-Neff period (panel C), the returns of January are relatively low compared to the rest of the year as shown by the difference between the January returns and the average for the rest of the year, which is -0.657 and not statistically different from zero (the t-statistic is only -0.74).

Although Neff's strategy did not involve necessarily shorting high P/E stocks, it is interesting to see whether the seasonality that we document in Table 5 above is applicable to the relative performance of low P/E versus high P/E stocks. Although we have seen that low P/E portfolios returns are significantly higher in the month of January than the rest of the year, we are now trying to document whether the cross-sectional difference between low P/E and high P/E stocks is also higher around the turn of the year. This would answer the question of whether the same type of seasonality is present in all P/E quintiles, or whether low P/E stocks are special in that sense. For that we calculate raw returns of zero-investment P/E portfolios, where we go long on low P/E stocks and short each one of the other P/E quintile, respectively. We present the results in Table 6.

The table presents the average monthly returns, in percentages, of a zero investment portfolio going long on low P/E stocks and short on high P/E stocks (Q1-Q5). The numbers presented in parentheses are t statistics testing whether the returns are significantly different from zero. The first column includes the whole period under study; the second and third columns isolate the Neff and non-Neff periods, respectively.

The results indicate that a zero investment portfolio makes the most sense in January (mostly during the Neff period). During the whole period other months that are of interest are July, September and March. They are statistically significant. The July results may be a result of the methodology that is used in the French database because the portfolios are refreshed in June, but the other months are a little anomalous and deserve further study.

Table 6: Differences in P/E Quintiles Portfolio Returns

	Whole Period (July 1951 - Dec 2009)	**Neff Period (July 1964 – Oct 1995)**	**Non-Neff Period (July 1951 – June 1964, Nov 1995 – Dec 2009)**
Mon	**Q1-Q5**	**Q1-Q5**	**Q1-Q5**
Jan	2.346	3.331	1.215
	(4.19)	*(3.97)*	*(1.81)*
Feb	0.578	0.459	0.715
	(1.12)	*(0.65)*	*(0.93)*
Mar	0.696	0.785	0.593
	(2.10)	*(1.72)*	*(1.21)*
Apr	0.678	0.429	0.964
	(1.72)	*(0.89)*	*(1.49)*
May	0.043	0.067	0.016
	(0.13)	*(0.13)*	*(0.03)*
June	0.239	0.071	0.431
	(0.61)	*(0.14)*	*(0.68)*
July	1.019	0.633	1.477
	(2.36)	*(1.11)*	*(2.24)*
Aug	0.363	0.721	-0.061
	(1.06)	*(1.47)*	*(-0.13)*
Sept	0.960	0.891	1.041
	(2.16)	*(1.60)*	*(1.44)*
Oct	-0.255	-1.254	0.930
	(-0.48)	*(-1.57)*	*(1.53)*
Nov	-0.341	-0.786	0.152
	(-0.82)	*(-1.40)*	*(0.25)*
Dec	0.620	0.345	0.925
	(1.42)	*(0.80)*	*(1.16)*

Conclusion

In analyzing some of the strategies of the time, like the Nifty Fifty, Blue Chip, and Growth, it is not surprising that John was able to grow Windsor into the largest mutual fund in America. Low P/E investing before his tenure, during his tenure and after his tenure was a superior investment style. We find evidence of an "incredible linearity" using P/E rank, which we believe justifies the use of P/E as an important yardstick for investors evaluating stocks. This type of linearity seems consistent with P/E being an important priced factor a la Ross's arbitrage pricing theory. As the first fund manager to press its advantage for the benefit of his mutual fund shareholders John rightfully deserves his place among the great investor of the 20th century. Low P/E investing in not inherently more risky than investing in glamour stock, but it gives far superior returns. The precise economic explanation for its superior results remains a mystery. During John's tenure the strategy got a lot of its punch at the turn of the year suggesting that a portfolio rebalancing explanation may play a role, but the January return effects dissipate when examining the return outside his tenure. Results presented here dictate continuing study of low P/E investing.

Notes

1. We take the inverse of the E/P ratio presented on French's website for consistency and ease of comparison with the literature. The earnings used in June of year t are total earnings before extraordinary items for the last fiscal year end (t-1). Price at time t is the total market value of equity calculated as price per share multiplied by shares outstanding at the end of December of t-1.
2. See website for detailed descriptions of portfolios formation.

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Exchange Rate Pass-through and Stock Market Development in Nigeria

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Abstract

Using time series data covering January 2003 to May 2010, this study examined the magnitude and response behaviour of exchange rate dynamics on stock prices in Nigeria. We utilised VAR methodology to test the impulse response and Chowbreak Point Tests for stability of our results across the selected business cycles in our sample. The results show a consistent negative significant relationship between exchange rate pass-through and stock market performance. The magnitude of this relationship was however inconsistent across the tested lags. This finding is puzzling and provides a strong reason for traders and investment advisers in the stock markets to pay more attention to the behaviour of exchange rate when making investment decisions.

Introduction

Stock and currency market share many puzzling empirical features, including excess volatility with respect to their fundamentals. Over the past few years, Nigerian economy has witnessed a spectacular increase in cross-border equity flows where expectations are affected by both equity returns in local currency and exchange rate fluctuation. Despite sharing common empirical features, little attention has been devoted to the analysis of these commonalities and the co-movement between these two asset prices.

Exchange rate volatility has implications for the financial system of a country especially the stock market. Interestingly, survey of available literature reveals divergent views of researchers on the issue of whether foreign exchange rate variability influences stock market volatility (Frank and Young, 1972; Solnik, 1987; Taylor and Tonks, 1989). Four events – Asian Currency Crises, the advent of floating exchange rate in the early 1970s, financial market reforms in the 1990s and the ongoing 2007-2010 financial crises prompted financial economists into re-examining the link between these two markets. Also, the internationalisation of capital markets has resulted in inflow of vast sums of funds between countries and in the cross listing of equities. This has therefore made investors and firms more interested in the volatility of exchange rate and its effect on stock market dynamics. Floating exchange rate appreciation reduces the competitiveness of export markets; and has a negative effect on the domestic stock market (Yucel and Kurt, 2003). But, for import dominated countries in West Africa, it may have a mixed effect on portfolio performance.

Among the few studies of exchange rate dynamics and stock markets on emerging markets are; Mishra (2004), Chortareas et al (2000); and Koutmoa et al (1993). Studies like Smith (1992), Solnik (1987), Aggarwal (1981), Frank and Young (1972), Phylaktis and Ravazzolo (2000), Granger et al. (2000), Abdalla and Murinde (1997), and Apte (2001) found a significant positive relationship between stock prices and exchange rates while others, such as Soenen and Hennigar (1998), Ajayi and Mougoue (1996), Mao and Kao (1990) have reported a significant negative relationship between the two variables. On the other hand, some studies, such as Bartov and Bodnar (1994), Frank and Young (1972), found very weak or no relationship between stock prices and exchange rates. On the issue of causation, most of the studies had mixed results (Morley and Pentecost (2000); Bahmani-Oskooee and Sohrabian (1992); Ibrahim (2000); Kanas (2000).

The openness of a country's economy is a key determinant of the behaviour of its market. More so, Nigeria is a classic example of an open economy with track record of huge bi-directional cross border capital flows. Most striking is the inverse response of the stock market relative to feedbacks from other developed and developing exchanges at the inception of the global financial crisis in 2007 and 2008. Therefore, the key objective of this study is to empirically examine the interaction between exchange rate dynamics and stock market performance in Nigeria. The sacred aspect of this study is our attempt to investigate the possibility of arbitrage while testing the depth of efficiency in the stock market using the behaviour of exchange rate over time.

The rest of the paper is organized into four sections. Section 2 contains trends of exchange rate and stock market indicators. Theoretical foundation and insights on the literature about stock market nexus with exchange rate are presented in section 3. The methodology, data and results are presented in section 4. Section 5 concludes with lessons for policy and suggestions for further studies.

Antecedent of Stock versus FX Market in Nigeria

The Nigerian Stock Market has experienced very robust growth over the last two decades in terms of activities, transaction types, volume and value of trade, number of listed companies and number of capital market operators. With the nation's transition to democracy in 1999 and the financial sector reforms since 2003, Nigeria's market took on a very strong up-tick which peaked in March 2008.

Table 1: Nigerian Stock Market Performance Indicators

Year	Exchange Rate (Year End)	NSE Index (Points)	Market Cap (N'bn)	Value Traded (N'm)	No. of Deals Annually	Listed Companies
2000	113.3	8,111	405.0	35,586.00	254,830	194
2001	112.3	10,963	580.5	63,837.00	419,834	194
2002	126.4	12,137	729.0	63,936.00	620,979	195
2003	136.5	20,128	811.5	116,232.00	720,979	200
2004	132.3	23,844	1,351.5	237,442.50	962,853	207
2005	128.5	24,085	2,079.0	270,876.00	1,016,989	212
2006	126.5	33,187	2,685.0	217,822.50	1,368,526	213
2007	116.3	57,990	12,678.0	2,597,299.50	2,537,215	212
2008	144.8	31,450	6,957.0	2,330,533.50	3,497,962	213
2009	148.4	20827	4,989.4	657,135.40	1,688,041	216
2010 (May)	148.9	25,398	6,177.9	383,199.31	1,066,582	214

Source: Nigerian Stock Exchange, CBN.

From the best performing equity market in 2007 of over 75% gain to one of the worst in 2008 and 2009 with a loss of 45% and 30% respectively. The first quarter of 2010 gained over 30% again returning it to one of the best performing markets in the world at the time. Nigerian market has once again exhibited its high volatility profile as optimistic analysts who projected about 50% return in the equity market got 20% loss.

On the other hand, exchange rate has been on the high side over the past decade. From the lows of N112/US$ in 2001 to the highs of over N160/US$ and N149/US$ in 2010. Economic reforms of 2003 to 2007 in Nigeria produced huge positive results for the equity market but its impact on exchange rate seemed to be minimal and short lived. We noted that local currency shed over 50% from its low of N117/US$ in 2007 to N160/US$ of 2008 when the equity market lost 45% at the same period of global economic and financial crisis. Interestingly, we are yet to see a repeat of this inverse co-movement for 2009 and 2010 as the gains in the equity market has only translated to a minimal fluctuating bound of +/-7% in exchange rate in the 18 months to May 2010.

Theoretical Foundation: Stock versus Forex Market

The parity theory provides economic explanation of the price at which two currencies should be exchanged, based on factors such as asset prices, inflation and interest rates. The theory suggests that when the parity condition does not hold, an arbitrage opportunity exists for market participants and this can impact on the asset prices depending the level of market efficiency.

The Asset Market Model holds that the inflow of money into a country by foreign investors for the purpose of purchasing assets such as stocks, bonds and other financial instruments can cause asset prices to rise. If a country is seeing large inflows by foreign investors, the price of its currency is expected to increase, as the domestic currency needs to be purchased by these foreign investors. This theory considers the capital account of the balance of trade and it has gained more acceptance as the capital accounts of countries are starting to greatly outpace the current account as international money flow increases.

From the economic point of view it is possible to trace a relationship between the stock market behavior and the exchange rate dynamics. In analysing the flows in the economies it is possible from economic theory to state that exchange rate performance does affect the real economy through international competitiveness and therefore on the balance of trade. The logical link between this and the company cash flows is immediate, just as it is to the stock prices. On the other hand economic theory also provides an inverted direction of causality, from the stock market to the exchange rate. This is because equities are by definition a very important fraction of wealth. Therefore swings in equity prices will affect the demand for money and finally the exchange rate determination.

The traditional models of the open economy have established the existence of a relationship between stock market performance and exchange rate behaviour. The models show that changes in exchange rates affect the competitiveness of firms as variations in exchange rate affect the value of the earnings as well as the cost of its funds because many companies borrow in foreign currencies to finance their operations and hence its stock price (Dornbusch and Fischer, 1980). An appreciation of the local currency, for example, makes exporting goods unattractive and leads to a decrease in foreign demand and hence revenue for the firm and its value would fall. This would also lead to a fall in the stock prices (Gavin, 1989).

Alternative approach to the study of the relationship between exchange rates and stock prices is provided by the portfolio balance models where the role of capital account transactions is stressed. For instance, a vibrant stock market would attract capital inflows from foreign investors, which increase the demand for its currency. The opposite would be the case with falling stock prices where the investors try to sell their stocks to avoid further losses and convert their money into foreign currency to move out of the country. Consequently, the local currency will depreciate. In the same vein, foreign investment in domestic equities could increase over time arising from the benefits of international diversification accruing to foreign investors. Thus, movements in stock prices may affect exchange rates and money demand because investors' wealth and liquidity demand could be a function of the performance of the stock market (Mishra, 2004).

Insights from Related Literatures

The relationship between exchange rates and stock market performance is of great interest to many academics and professionals, since they play a crucial role in the economy. Nonetheless, results are somewhat mixed as to whether stock indexes lead exchange rates or vice versa and whether feedback effects (bi-causality) even exist among these financial variables.

Aggarwal (1981) argued that changes in exchange rates provoke profits or losses in the balance sheet of multinational firms, which induces their stock prices to change. In this case, exchange rates cause changes in stock prices (traditional approach). Dornbusch (1975) and Boyer (1977) presented models suggesting that changes in stock prices and exchange rates are related by capital movements. Decrease in stock prices reduces domestic wealth, lowering the demand for money and interest rates, inducing capital outflows and currency depreciation. Bahmani-Oskooee and Sohrabian (1992) analysed the relationship between stock prices and exchange rates in the US economy. They found no long-run relationship among these variables, but a dual causal relationship in the short-run using Granger (1969) causality tests. Using S&P 500, effective exchange rate and monthly data over the period, July 1973 to December 1994, Bartov and Bodnar (1994) found that lagged, and not contemporaneous changes in US dollar exchange rates, explain firms current stock returns.

Ratner (1993) applied co-integration analysis to test whether US dollar exchange rates affect US stock prices, using monthly data from March 1973 to December 1989. His results indicated that the underlying long-term stochastic properties of the US stock index and foreign exchange rates are not related, since the null of no co-integration could not be rejected, even when dividing the sample into sub-periods.

Ajayi and Mougoué (1996) analysed the relationship between stock prices and exchange rates in eight advanced economies (Canada, France, Germany, Italy, Japan, the Netherlands, the United Kingdom and the United States). Using an error correction model, they found significant short and long run feedback between these two variables. Abdalla and Murinde (1997) investigated interactions between exchange rates and stock performance in India, Korea, Pakistan, and the Philippines using monthly observations in the period from January 1985 to July 1994. Within an error correction model framework, they found evidence of unidirectional causality from exchange rates to stock prices in all countries, except for the Philippines. There, they found that stock prices influence exchange rates.

Ong and Izan (1999) used weekly data of "spot and 90-day forward" exchange rates for Australia and the G-7 countries and "spot and 90-day forward" futures prices for equity prices in Australia, Britain, France and the US, during the period from October 1986 to December 1992. They were unable to find a significant relationship between equity and exchange rate markets. They suggested that the use of daily data (or even intra-day) data could improve their empirical results.

Ajayi et al (1998) used daily data and reported that causality runs from the stock market to the currency market in Indonesia and the Philippines, while in Korea it runs in the opposite direction. No significant causal relation is observed in Hong Kong, Singapore, Thailand, or Malaysia. However, in Taiwan, they detected bi-directional causality or feedback. Furthermore, contemporaneous adjustments are significant in only three of these eight countries. In developed countries, they found significant unidirectional causality from stock to currency markets and significant contemporaneous effects.

Granger et al. (2000) found strong feedback relations between Hong Kong, Malaysia, Thailand and Taiwan. They used daily data and their sample period started January 3, 1986 and finished June 16, 1998. Furthermore, they found that the results are in line with the traditional approach in Korea, while they agreed with the portfolio approach in the Philippines. Nieh and Lee (2001) found no significant long-run relationship between stock prices and exchange rates in G-7 countries, using both the Engle-Granger and Johansen's co-integration tests. Furthermore, they found ambiguous and significant, short-run relationships for these countries. Nonetheless, in some countries, both stock indexes and exchange rates may serve to forecast the future paths of these variables. For example, they found that currency depreciation stimulates Canadian and UK stock markets with a one-day lag, and that increase in stock prices cause currency depreciation in Italy and Japan, again with a one-day lag.

In general, empirical findings suggest that there are no long-run equilibrium relationships between these two financial variables (exchange rates and stock market performance) in most countries. However, many studies have found that these variables have "predictive ability" for each other, although the direction of causality seems to depend on specific characteristics of the country analysed. To the best of my knowledge, this is the first study that seeks to address the pass-through effect of exchange rate on stock market performance in a West African Economy.

Data and Models

Data

This Study is aimed at examining the interaction between exchange rate dynamics and key stock market indicators in Nigeria. Time series data on exchange rate, stock exchange index were used. Exchange rate was sourced from CBN and stock exchange index was sourced from the Nigerian Stock Exchange (NSE). We used daily data because it is arguably more revealing given that the use of monthly data may not be adequate to capture the effects of short-term FX rate and stock price movement due to high volatility. NSE index is the key performance indicator becouse it incorporates the price movement of all the equities listed on the NSE at any point in time. The data series covers two business and political cycles from January 2003 to May 2010.

Models

Vector Autoregression (VAR)

The paper employs vector autoregression (VAR) methodology to study the magnitude of effect and the response to impulse function of exchange rate with respect to stock prices. Vector autoregression (VAR) is an econometric model used to capture the evolution and the interdependencies between multiple time series, generalizing the univariate AR models. All the variables in a VAR are treated symmetrically by including for each variable an equation explaining its evolution based on its own lags and the lags of all the other variables in the model.

Given a VAR of order p, where the order p represents the number of lags, that includes k variables, VAR model can be expressed as:

$$x = T_0 + \sum_{i=1}^{p} T_i \theta_i x_{t-i} + u_i \qquad (1)$$

Where xt = [x1t ... xkt]' is a column vector of observation on the current values of all variables in the model, Ti is k×k matrix of unknown coefficients, T0 is a column vector of deterministic constant terms, ui is a column vector of errors with the properties of

$$E(u_i) = 0 \text{ for all t}, E(u_s u_t) = \Omega \text{ if } s=t \quad \text{and } E(u_s u_t) = 0 \text{ if } s \neq t \qquad (2)$$

Where Ω is the variance-covariance matrix. ut's are not serially correlated but may be contemporaneously correlated. Thus, Ω is assumed to have non-zero off-diagonal elements.

Impulse Response Function

In a VAR system, the examination of the estimated coefficients on successive lags has no sufficient information on the dynamic relationships among the variables in the system. Rather, it is useful to trace out the system's response to typical random shocks that represent positive residuals of one standard deviation unit in each equation in the system. Therefore, Sims (1980) proposes the use of impulse response and variance decomposition to assist in achieving this logical interpretation of the VAR system.

Assuming a 2-variables VAR (1) model specified as

$$\begin{bmatrix} x_{1t} \\ x_{2t} \end{bmatrix} = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix} \begin{bmatrix} x_{1t-1} \\ x_{2t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \tag{3}$$

A disturbance in ε_{1t} has an instant and direct effect on x_{1t}. In period t+1, that disturbance in x_{1t-1} affects x_{2t-1} through the first equation and also affects x_{2t-1} through the second equation. These effects work through to period t+2, and so on. Thus, a random shock in one innovation in the VAR sets up a chain reaction over time in all variables in the VAR. Impulse response functions calculates these chain reactions.

Impulse response functions is confronted with one limitation; that is, a disturbance in one innovation is not contemporaneously isolated from the other innovations in the system, although it ultimately leads to a chain reaction over time in all variables in the system. It is doubtful from the above bivariate model to hypothesize that one innovation receives a disturbance while the other does not. A solution to this problem is achieved by transforming the innovations to produce a new set of orthogonal innovations, which are pair wise uncorrelated and have unit variance.

Chow Breakpoint Tests

We also apply the Chow breakpoint test to strengthen reliance on our results by dividing our sample period into different reproduction periods (business cycles). In econometrics, the Chow test is most commonly used in time series analysis to test for the presence of a structural break by splitting the original data. By this we are able to ascertain by how much the coefficients of our least squares estimates vary from one period to the other.

Empirical Analysis and Results

Unit Root Test

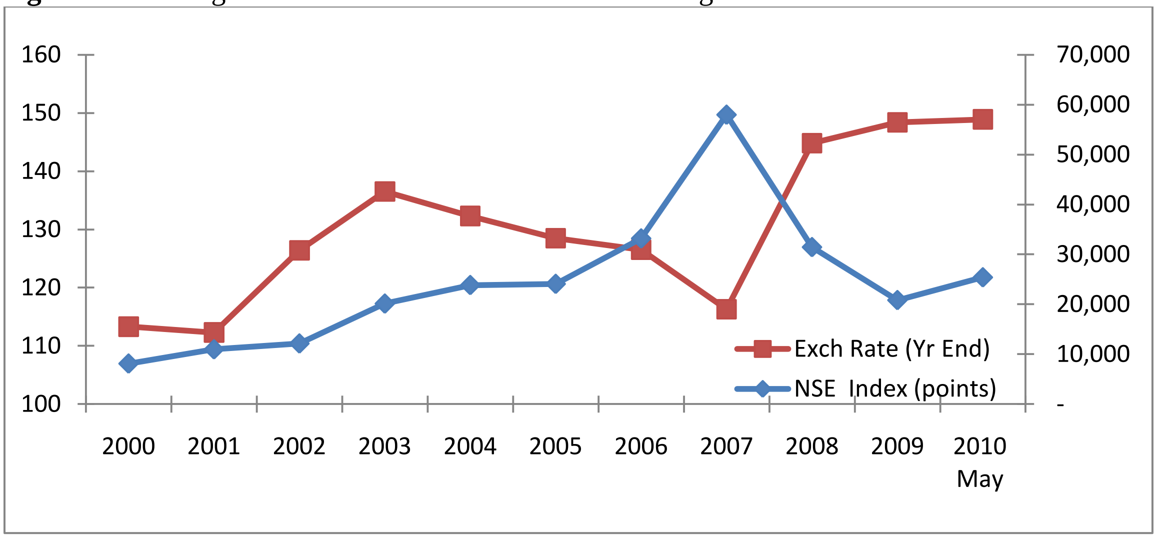
The paper conducts unit root tests of the variables with Phillips Perron (PP) and Augmented Dickey Fuller (ADF). Outcomes of the tests are presented in Table below. According to the PP test, all variables in log level (FXrate and NSEindex) rejected the null hypothesis at all the significant levels with a constant and a trend. ADF tests also collaborated the same result at all the significant levels by rejecting the null hypothesis that the variables are level and trend stationary. This suggests that the variables have a strong attribute of sustaining a long run relationship. What we cannot say for sure at this stage of our estimate is the feature of their relationship and inherent lag determinants.

Magnitude and Significance of Exchange Rate Pass-through Effects

According to past studies, working with daily data, it is expected that 30 days lags variables should be enough to conduct the VAR estimation for our sample size without problems. However, we avoided the optimal lag length problem by stretching our VAR estimates up to 245 days (maximum trading days in a year). We reported results up to 30 day lag in the apendix due to lack of space. Estimated results for all the lag lengths are available on request. Interestingly, we find a consistent negative significant relationship between exchange rate pass-through and stock market performance. The magnitude of this relationship was however inconsistent across the tested lags. We are particularly excited to discover that over a distant lag of up to 14 trading days, the significant level of negative relationship tends to become stronger and this momentum relaxes as the estimate move towards the 30th trading day.

In a more rousted form, we tested for biavariate relationship between market performance and exchange rate trends using VAR estimates. Contrary to the popular view that exchange rate movement in Nigeria is better predicted using crude oil prices, we find that negative fluctuation of stock prices might be another strong catalyst for exchange rate volatility. This hypothesis became evident when we endogenize the NSE index in our VAR estimates over the tested lags. We remind ourselves of the spectacular case in 2008 trading year when the Nigerian stock market witnessed a major flight to safety of foreign portfolio holders and institutional investors who accounted for 7% and 70% of the market respectively at the time. Negative fluctuation of stock prices coupled with the wave of global economic crisis created a massive demand for FX making the local currency to see an all time high depreciation of 55% in 30 days. This trend surpassed the speed of devaluation during the Structural adjustment Programme (SAP) of the 1980's. All in all, our results confirm a more powerful negative response flowing from stock market losses to exchange rate movements.

Figure 1: Exchange Rate and Stock Index Interaction in Nigeria



Chow Breakpoint Stability Test

To strengthen reliance on our results, we subjected our estimates to Chow Breakpoint stability test by dividing our sample period into different reproduction periods (business cycles). We are happy to note that our VAR and Ordinary Least Square (OLS) results are stable and consistent over all the tested reproductions period as the F-statistics and probability show high insignificant levels.

Table 2: Selected Breakpoints (Business Cycles)

Selected Break Points	Economic/Political Landmarks
May 2004	Kick off of CBN banking sector consolidation and reform of the financial sector under President Olusegun Obasanjo.
May 2007	Commencement of a new political administration under President Umaru Musa Yar'Adua/ Goodluck Jonathan and the kick off of remarkable reversal of completed and ongoing economic reform programs.
November 2008	The peak of exchange rate rise during the global economic and financial crisis and exit of institutional/foreign portfolio holders from the Nigerian Stock Market.
September 2009	The climax of recent banking sector crisis and CBN's bail-out (N620 billion) of the sector that account for over 50% of Stock market capitalisation in Nigeria at the time.

Source: Author's compilation

Impulse Response Functions

According to Sims (1980), most estimated coefficients from a VAR model are not statistically significant. Therefore, the impulse response functions and variance decompositions are further examined. Impulse response functions are dynamic simulations showing the response of an endogenous variable over time to a given shock, while variance decompositions show the contributions of each source of shocks to the variance of the future forecast error for each endogenous variable.

Thus, attempt is made to examine the effect of exchange rate shock on real stock returns using impulse response functions. The figures show that in all the sub-samples (2004-2007; 2007-2008 and 2008-2009), an exchange rate shock has a negative and statistically significant impact on real stock returns at all the significant levels. Outcomes have different magnitudes depending on the length of our prediction lag.

Conclusion and Policy Lessons

Cross country studies over the years revealed that FX and stock market variables have "predictive ability" for each other. However, the direction and speed of response seems to depend on specific characteristics of the country analysed. This study estimates the effects of exchange rate pass-through on the stock returns in Nigeria from January 2003 to May 2010 using a bivariate VAR analysis.

Empirical results show a consistent negative significant relationship between exchange rate pass-through and stock market performance. The magnitude of this relationship was however inconsistent across the tested lags. We are particularly excited to discover that over a distant lag of up to 14 trading days, the significant level of negative relationship tend to become stronger and this momentum relaxes as the estimate move towards the 30th trading day.

Contrary to the popular view that exchange rate movement in Nigeria is better predicted using crude oil prices, we find that negative fluctuation of stock prices might be another strong catalyst for exchange rate volatility. This hypothesis became evident when we endogenise the NSE index in our VAR estimates over the tested lags. We confirm that a more powerful negative response flow from stock market losses to exchange rate movements.

To strengthen reliance on our results, we subjected our estimates to Chow Breakpoint stability test by dividing our sample period into different reproduction periods (business cycles) based on economic and political landmarks over the sample period. Incidentally, that our VAR and Ordinary Least Square (OLS) results are stable and consistent over all the tested reproductions periods.

The study has not been able to address the following issues. First, to analyze the asymmetric effect of exchange rate pass through in oil-exporting and oil-importing countries by confirming whether or not exchange rate changes have the asymmetric effect on the stock market. Also, there is need to examine the industrial classification of firms most affected by exchange rate pass through and see whether or not the results differ from industry to industry. Finally, there is need to present an economic model relating exchange rate fluctuations to firm's dividends and performance. These issues are subjects for future work.

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Appendix

Figure 2: FDI Flows in Nigeria

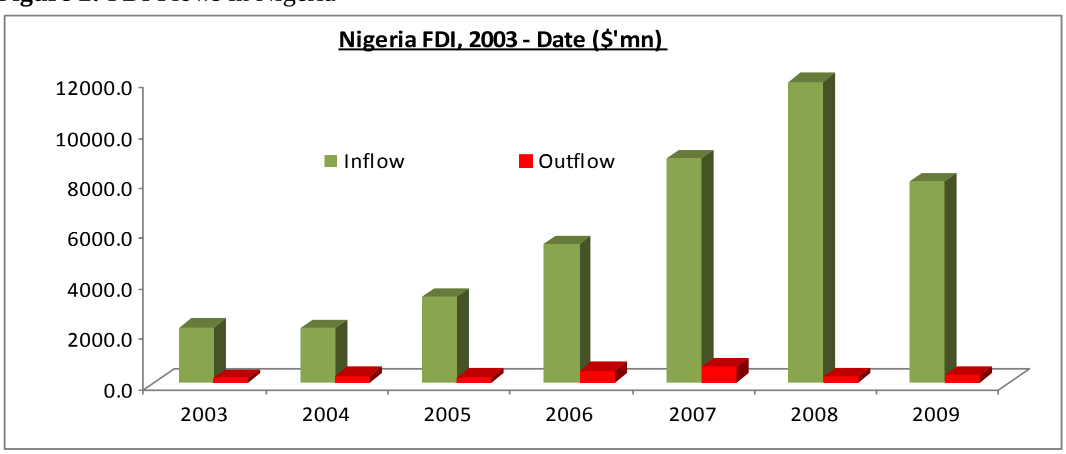

Table 3: Classification of FDI, 1990 to 2009

Year	Inward investment ($'m)				Outward investment ($'m)			
	Stocks	Reinvested Earnings	Intra-Company Loans	Total Inflow	Stocks	Re-invested Earnings	Intra-Company Loans	Total Outflow
1998	91.5	479	639.8	1,210.10	34	50.2	74	158.8
1999	5.2	412	760.5	1,177.70	1.1	48	123.7	172.8
2000	1.6	433	875.4	1,309.70	0.2	40	128.8	168.9
2001	4.1	1,088	185	1,277.40	4.6	50	39.3	93.9
2002	7.8	1,704	328	2,040.20	1.4	77.8	93	172.2
2003	10.2	1,873	288	2,171.40	0.1	68.3	98.9	167.3
2004	18.9	1,708	400.7	2,127.10	0.1	99.8	160.8	260.8
2005	30.2	2,732	641	3,403.30	0.1	70.1	129.9	200.1
2006	49.1	4,440	1041.7	5,530.40	0.2	169.9	290.7	460.9
2007	130	7,172	1682.7	8,984.50	0.3	238.2	389.6	628.2
2008	840			12,000.0				250
2009*				6,232.3				

Source: United Nation Trade and Development (UNTAD)

Table 4: Share of Foreign Portfolio on the Nigerian Stock Market

Year	Value of Shares Traded (N'bn)	Contribution of Foreign Equity Flow (N'bn)	Foreign Equity Flow as % of Market Turnover
2004	**220.01**	**6.60**	**3**
2005	255.03	**7.65**	**3**
2006	464.27	23.21	5
2007	2,077.80	166.22	8
2008	2,330.53	163.14	7
2009	630.85	37.85	6

Source: NSE, Author's Calculation

Table 5: Chow Breakpoint Tests

Chow Breakpoint Test: 5/29/2004			
F-statistic	524.6954	Probability	0.000000
Log likelihood ratio	470.3642	Probability	0.000000

Chow Breakpoint Test: 5/29/2007			
F-statistic	246.6582	Probability	0.000000
Log likelihood ratio	233.8344	Probability	0.000000

Chow Breakpoint Test: 11/03/2008			
F-statistic	174.7304	Probability	0.000000
Log likelihood ratio	168.2069	Probability	0.000000

Chow Breakpoint Test: 8/28/2009			
F-statistic	207.0359	Probability	0.000000
Log likelihood ratio	197.9282	Probability	0.000000

Table 6: Vector Autoregression Estimates

FX Rate impacts on Stock Prices		**Stock Price Impacts on FX Rate**	
Sample(adjusted): 2/13/2002 - 4/29/2010		Sample(adjusted): 2/13/2002 - 4 /29/2010	
Included observations: 2142		Included observations: 2142	
after adjusting endpoints		after adjusting endpoints	
	NSEINDEX		**FXRATE**
NSEINDEX(-1)	1.124813	FXRATE(-1)	0.998904
	(0.02177)		(0.02177)
	[51.6773]		[45.8847]
NSEINDEX(-2)	0.077946	FXRATE(-2)	0.001419
	(0.03276)		(0.03077)
	[-2.37929]		[0.04610]
NSEINDEX(-3)	-0.038115	FXRATE(-3)	-0.000730
	(0.03280)		(0.03077)
	[-1.16192]		[0.02372]
NSEINDEX(-6)	-0.007381	FXRATE(-6)	-0.002439
	(0.03282)		(0.03077)
	[0.22492]		[-0.07927]
NSEINDEX(-14)	-0.020422	FXRATE(-14)	-0.000160
	(0.03281)		(0.03077)
	[0.62248]		[-0.00520]
NSEINDEX(-15)	-0.012518	FXRATE(-15)	-0.000139
	(0.03281)		(0.03077)
	[-0.38154]		[-0.00452]
NSEINDEX(-21)	-0.008477	FXRATE(-21)	-0.003152
	(0.03281)		(0.03077)
	[-0.25839]		[-0.10243]
NSEINDEX(-25)	-0.005568	FXRATE(-25)	-0.001263
	(0.03281)		(0.03077)
	[-0.16969]		[0.04105]
NSEINDEX(-29)	-0.017862	FXRATE(-29)	-0.000335
	(0.03276)		(0.03077)
	[-0.54526]		[-0.01090]
NSEINDEX(-30)	0.015883	FXRATE(-30)	-0.001489
	(0.02175)		(0.02190)
	[0.73037]		[-0.06802]
C	-200.0629	C	-0.203610
	(72.2869)		(0.38837)
	[-2.76762]		[0.52427]
FXRATE	1.947889	NSEINDEX	9.07E-06
	(0.56502)		(4.9E-06)
	[3.44747]		[1.86065]
R-squared	0.998015	R-squared	0.986610
Adj. R-squared	0.997986	Adj. R-squared	0.986413
Sum sq. resids	9.41E+08	Sum sq. resids	22303.53
S.E. equation	667.6980	S.E. equation	3.251214
F-statistic	34218.04	F-statistic	5015.122
Log likelihood	-16954.46	Log likelihood	-5548.725
Akaike AIC	15.86038	Akaike AIC	5.210761
Schwarz SC	15.94507	Schwarz SC	5.295459
Mean dependent	26992.13	Mean dependent	124.2952
S.D. dependent	14876.87	S.D. dependent	27.89241

Note: Standard errors in ( ) & t-statistics in [ ]

New Information Regarding Consumption and Wealth Asymmetries

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Abstract

This paper examines the existence of asymmetric behavior in the consumption, wealth, and income relationship. This relationship is modeled using a vector error correction methodology, which is common in the recent literature. Unlike previous research, we find mixed cointegration results. Consumption asymmetry is a common finding in previous research, but the findings here reject long run asymmetric behavior. Further, wealth is shown to respond symmetrically to shocks in the long run. Therefore, the results here support those of Lettau and Ludvigson (2004), since wealth is the variable that responds most dramatically to shocks in the long run.

Introduction

Much of the empirical literature surrounding the linkages between wealth and consumption starts with Modigliani's 1971 paper, where he finds that consumption increases about five cents for every dollar of new wealth. Prior to this, Ando and Modigliani (1963) actually develop the Life-Cycle model hypothesis. Even in 1963, the first attempt to estimate the magnitude of the effect provides an estimate in the 4 cents to 9 cents range. However, Ando and Modigliani results span the timeframe of 1929 through 1959, excluding the years 1941 - 1946. In the last five years of Ando and Modigliani's data, the change in household wealth ranged from 34.4B to 93.5B. Contrast that with the last five years; from 2005 through 2009, changes in household wealth ranged from -6,700B (in 2005) to. 12,700B (in 2008). Moreover, the first decade in this new millennia has seen the first ten year span where appreciation in share prices is essentially flat, combining share price appreciation in the early years of the decade with several recent years of severe price declines. Couple this with the unprecedented collapse in housing values experienced between 2006 and the present, and we now have an opportunity to observe interesting changes in the pattern of consumption.

In 2004, Lettau and Ludvigson confirm that about 5 cents of every additional dollar of wealth is consumed; however, they argue that wealth must be separated into transitory and permanent shocks. The 5 cents figure that they estimated is for *permanent* changes in wealth. In their 2004 paper, Lettau and Ludvigson find that the majority of changes in wealth are temporary, thus the 5 cent figure may overstate the true effect of wealth changes on consumption. The underlying premise behind the Life-Cycle hypothesis, and the similar Permanent Income hypothesis (Friedman 1957), is brilliant in its simplicity; individuals adjust their consumptive patterns to changes in income. If we accept that each hypothesis is true, then we would expect that positive, epsilon changes in wealth (or permanent wealth) drive increases in consumption; likewise, a negative, epsilon change in wealth should lead to a decrease in consumption of equal magnitude. Any variation from this expected pattern would indicate the presence of an asymmetric relationship between wealth and consumption. Much of the literature that attempts to explain this asymmetric response centers around the most likely possibility that credit constraints hamper an individual's ability to smooth life time income. What follows is a review of the literature, a description of the data, our estimations and analysis, and a conclusion.

Previous Literature

Patterson (1993) develops a model that explains how credit constrained customers may still respond to changes in interest rates. Credit constraints have emerged as the most likely source of consumption asymmetries, over and above the other likely suggestion, which is myopic behavior among consumers. While Patterson's (1993) model is theoretical, he notes that simply including changes in income are likely to be insufficient as a remedy, as it will not capture the "likely asymmetry between the response of consumption to increases and decreases in income" (pg 406). Shea (1995) uses aggregate time series data to test the two competing explanations offered as a reason why the empirical evidence fails to confirm the life cycle or permanent income hypothesis when aggregate data is used. The two prevailing explanations, either myopia or liquidity constraints, are both testable using aggregate time series data. Shea finds that neither explanation is statistically valid. He estimates the change in consumption as a function of positive changes in expected income and negative changes in expected income, controlling for the real interest rate. Shea finds that aggregate consumption responds more to negative changes in predictable income than to positive changes, which is contrary to the life cycle/permanent income hypothesis, myopia or liquidity constraints.

Garcia, Lusardi and Ng (1997) use a switching regression model in a panel approach with Consumer Expenditure Survey data to again test liquidity constraints and myopia as explanations for consumption asymmetries. They use a logit model to estimate the probability that a consumer is liquidity constrained, and then test the liquidity constrained consumer's consumption responses to changes in income. They find that liquidity constrained consumers responded to the constraint, but not asymmetrically. They do not find evidence to support the myopic explanation. Noteworthy is the fact that unconstrained consumer's responded to negative income changes differently than positive income changes.

Mehra (2001) looks at the effects of wealth on consumption using an error correction model. Prior to Mehra, much of the research corrected for potential spurious regression problems by estimating in first differences. Mehra uses the Phillips-Oularis residual-based cointegration test and confirms that real consumer spending, labor income, and nonequity net worth are all cointegrated, after discovering unit roots in the presence of a linear trend. While estimated differently, Mehra still finds marginal propensity to consume from wealth of 3 cents. Mehra's estimates are in line with those suggested by Poterba (2000). The most intuitively and parsimonious calculations in Poterba (2000) take a $1 increase in wealth and solve the annuity problem for varying after-tax returns and life spans, suggesting that marginal propensity to consumer from wealth should range from 3 cents to 10 cents, assuming no bequest motive.

Lettau and Ludvigson (2004) attempt to disaggregate consumption, wealth, and labor income into their respective permanent and transitory components. Using the log of asset wealth, as defined by real per capita household net worth, the log of real, per capita expenditures on nondurables and services, and the log of after-tax real per capita labor income[1], Lettau and Ludvigson find that all variables are first-difference stationary and move together in the long-run through a single, cointegrated relationship. Using variance-decomposition analysis, where the shocks are separated into permanent and transitory shocks, Lettau and Ludvigson show that consumption responds strongly to permanent changes in wealth and labor income, while barely responding to transitory changes in the respective variables. This is very interesting because the variance decomposition analysis shows that the growth of wealth responds almost entirely (88%) to transitory shocks. As Lettau and Ludvigson report, the estimated wealth effect of four to five cents is most probably overstated, as most of the variance of wealth is transitory, and consumption barely responds to transitory changes in wealth.

Apergis and Miller (2006) attempt to model both short-run and long-run asymmetries in a single equation framework. Using real personal consumption per capita, real after-tax nominal labor income per capita, the consumer price index, and real stock market capitalization per capita, from 1957 through 2002, Apergis and Miller find all variables but CPI are first-difference stationary. Using three lags, a single cointegrated relationship is estimated, which is consistent with previous estimations in the literature. In the long run, Apergis and Miller estimate the elasticity of real per capita consumption as a function of real per capita stock market value is .375%. Through a series of F-tests, testing for asymmetries using a modified error-correction model estimated with 2 lags, the authors find that negative stock market values have a larger effect than positive stock market values. As they note, their finding is consistent with Kahnemen and Tversky's theory of loss aversion.

Treeck (2008) examines the issue following the framework in Apergis and Miller (2006), trying to model both short-run and long-run asymmetries in the relationship. Treeck models consumption in an asymmetric error-correction method, where changes in consumption are a function of lagged consumption, income and wealth, as well as lagged values of changes in the respective variables, with wealth subdivided into positive and negative changes. Results in the working paper indicate that in the long run, consumption responds asymmetrically to changes in labor income, while there is no statistical evidence of asymmetric responses to wealth.

Gabriel, Alexandre and Bacao (2008) model the consumption-wealth ratio using a Markov-switching-vector error correction model. Gabriel *et al.* pursue the Markov-switching methodology after discovering issues with the model assumptions used by Lettau and Ludvigson (2001, 2004). While Gabriel *et al.* are quick to point out that the conclusions from Lettau and Ludvigson are unchanged, they pursue the nonlinear, Markov-switching method of estimation to correct for these issues. Using this Markov-switching framework, Gabriel *et al* show that during periods of turbulence, wealth provides the adjustment to return to long-run equilibrium; however, when the economy is experiencing tranquility, then it is consumption that adjusts more. Curiously, one assumption underlying the model is that consumers who accurately perceive the nature of the shock, be it temporary or permanent, react as the life cycle or permanent income hypothesis would suggest; for shocks perceived incorrectly, there is a one-period mistake in their response, because they misperceive the true nature of the shock, but they correct the error in the second period.

Bostic, Gabriel and Painter (2009) approach the consumption-wealth linkage from a unique perspective by using micro level data. They use a unique process to match entries from the Survey of Consumer Finances conducted by the Federal Reserve to entries from the Consumer Expenditure Survey conducted by the Bureau of Labor Statistics. Modeling total consumption, as well as durable goods consumption, as a function of income, financial wealth, housing values, and other real estate holdings (admittedly a very small category), they find elasticities of financial wealth and housing wealth of .02 and .05, respectively. More germane to the investigation of asymmetric consumption responses, Bostic *et al.* attempt to assess the effect of credit constraints on households. As we know, if asymmetric behavior is revealed in aggregate data, it must begin with the household. Using a dummy variable to identify credit-constrained households (households with a loan-to-value ratio that exceeds 90%) and an interaction term between the dummy variable and house price, the authors do not find a statistically significant link between total consumption and either variable. Re-examining the dataset, including only households who self-identified as credit constrained, there is still no statistical connection.

Chen, Chen and Chou (2010) investigate the relationship between housing wealth and consumption in Taiwan. Using several different models, they find that durable consumption is affected by changes in housing wealth. Interestingly, the authors did not expect to find a link between nondurable consumption and housing wealth, as home equity loans "are nearly nonexistent in Singapore as well as Taiwan (pg 9 of preliminary draft)." Using an error correction model, they find that both disposable income and changes in real house prices produce asymmetric consumption responses in Taiwan. This confirms results from an earlier model they estimated using a threshold regression and an instrumental variable approach. These results only hold when consumers are credit-constrained. As with the earlier literature, liquidity or credit-constraints are the explanatory factor for deviations from the permanent income hypothesis.

Odusami (2010) has an interesting twist to the relationship between consumption and aggregate wealth. He posits that the short-run changes in the consumption-aggregate wealth ratio can be predicted by movements in the price of crude oil. Using a multivariate threshold vector autoregressive (MTAR) model, where the log of the consumption-wealth ratio is explained by one past observation of itself, as well as an endogenously determined number of past observations of the quarterly change in the real price of oil, Odusami (2010) finds that changes in the price of oil have a greater affect on transitory changes in the consumption-wealth ratio during the period 1959 - 1981, then they do during the period of 1981 – 2007.[2] Odusami compares the performance of the MTAR model to a similar class of models, the self-exciting threshold autoregressive model (SETAR), but fails to find the SETAR model superior.

Ibrahim and Habibullah (2010) consider the relationship between stock market wealth and aggregate consumption in Malaysia. Following Enders and Siklos (2001) in their approach, they model aggregate consumption as a function of real income and real share prices, testing for asymmetric cointegration in this relationship. Using the standard Engle-Granger (1987) approach to cointegration, they do not find statistical evidence of a long-run relationship, nor do they find a long-run relationship when modeled through the Enders and Siklos methods. However, when they consider the ratio of consumption to income, as well as wealth to income, asymmetric cointegration results confirm a long-run relationship. The authors do not find evidence of adjustment when consumption-income is below its long run value; however, this may be a function of the model, specifically that wealth is proxied by increases in share prices as represented by the Kuala Lampur Composite Index.

We test for cointegration in the relationship between U.S. aggregate consumption, labor income, and wealth, without assuming asymmetries, all-the-while noting that Balke and Fomby (1997) show that standard cointegration tests may fail if asymmetries are present. Thus, we extend our work to follow Enders and Siklos (2001) to test for asymmetric cointegration. Our work becomes particularly important as we have now experienced substantial declines in housing wealth, as well as an entire decade where stock returns, as measured by the Standard and Poors 500, have been virtually flat. Using and updating the data set by Lettau and Ludvigson (2004) through 2010, we have a sufficiently long dataset that allows us to model both short-run and long-run asymmetries in the relationship. To begin with, we only examine the long-run relationship.

Data

This research uses consumption, labor income, and wealth data, which is common in the literature. The consumption and labor income measures are calculated from data provided by the Bureau of Economic Research (BEA). The consumption measure included here is the sum of nondurable and services consumption excluding clothing and footwear. Specifically, the consumption measure comes from table 2.3.5 of the National Income and Product Accounts (NIPA). Labor income is calculated using the data provide in NIPA table 2.1 "Personal Income and Its Disposition". Labor income is the sum of wages and salaries, transfer payments, and other labor income less social insurance contributions and taxes.[3] Finally, the wealth data is available from the Flow of Funds Accounts, which is provided by the Board of Governors of the Federal Reserve System. Table B.100 provides wealth of households and nonprofit organizations. Wealth is measured as household assets minus household liabilities. All series are converted into real terms using the personal consumption expenditures price index provided in NIPA table 2.3.4 at the BEA. Further, the data is adjusted to per capita values using population and are included in natural log levels. The data is provided at a quarterly frequency and ranges from the first quarter of 1952 through the first quarter of 2010.

Unit Root Tests

Unit root tests on consumption, labor income, and wealth using equation (1) show that all variables appear non-stationary in natural log levels, since testing fails to reject the null hypothesis of a unit root. Unit root testing rejects the null of non-stationarity in first-differences of natural logs. A series is non-stationary (i.e. has a unit root) if testing fails to reject the null hypothesis that γ is statistically zero.

(1)

Symmetric and Asymmetric Cointegration Tests

We test for cointegration between the three variables using two common tests: the Johansen test (1988) and the Engle-Granger test (1987). We estimate the cointegrating relationship in the Engle-Granger test using the Dynamic OLS (DOLS) method of Stock and Watson (1993). Table 2 presents the results of the two cointegration tests. Interestingly, both tests fail to reject the null of no cointegration between the three variables.[4]

Table 1: Cointegration Tests

Panel A: Johansen Test					
Hypothesized number of Cointegrating Vectors	Wealth Parameter	Labor Income Parameter	Constant	Trace Test Statistic	5% Critical Value
0	0.450	0.640	1.699	23.921	29.797
Panel B: Dynamic OLS (Engle-Granger Test)					
	Wealth Parameter	Labor Income Parameter	Constant	Test Statistic	5% Critical Value
	0.392	0.706	-1.657	-2.848	-3.760

Two lags used in the Johansen test, as suggest by Akaike Criterion and Schwarz Criterion. Four leads and lags of first-differences included in the DOLS cointgration equation.

Given the result of no cointegration, we next examine the issue of cointegration under asymmetric adjustment. The common tests used here assume symmetric adjustment and are misspecified if the adjustment is asymmetric. Enders and Siklos (2001) suggest that these cointegration tests have low power in the presence of asymmetric adjustment. To test for cointegration in the presence of asymmetric adjustment, we use the methods presented by Enders and Siklos. Specifically, we test for cointegration using two alternative forms of error-correction: a threshold autoregressive model (TAR) and a momentum threshold autoregressive model (MTAR).

Specifically, we test for cointegration in the presence of asymmetry around the equilibrium (cointegrating relationship) defined under the Johansen test results in Panel A of Table 1 and under the Dynamic OLS test results in Panel B of Table 1. The TAR test is specified in equation (2), which tests for asymmetric adjustment relative to the equilibrium derived in Table 1. The MTAR model is specified in equation (3). Under the MTAR specification, the adjustment of the variables is allowed to vary according to the previous period's change in the error-correction term. In both tests, the attractor τ is set equal to zero. As mentioned, the variable *e* in equations (2) and (3) is the error correction term derived from the respective cointegration test.

(2)

In the TAR model, the Heaviside indicator, I_t, is determined by the value of the error-correction term relative to equilibrium. Starting at equilibrium, an increase in consumption or a decrease in wealth, holding all else constant, produces a positive value of *e*, and I_t equals one. In the MTAR model, however, the Heaviside indicator is determined by the previous period's change in the error-correction term. For example, an increase in consumption that is relatively larger than a given increase in wealth produces a positive value of *e* and the Heaviside indication equals one.

(3)

There are two tests and critical values provided in Enders and Siklos to test for cointegration in the presence of asymmetric adjustment, the t-max test and the Φ test. The t-max test uses the maximum t-value under the null hypothesis of ρ_i=0. Enders and Siklos note that ρ_1<0 and ρ_2<0 are necessary conditions for convergence, which is tested under the t-max statistic. The Φ statistic is an F-test of the joint hypothesis $\rho_1=\rho_2=0$. The Φ test is a test of cointegration in the presence of asymmetry. Results of the TAR and MTAR test are given in Table 2. Columns (1) and (2) provide coefficient estimates and t-statistics for ρ_1 and ρ_2. Column (3) gives the F-statistic for the null hypothesis of $\rho_1=\rho_2=0$, and the test of symmetry ($\rho_1=\rho_2$) is provided in column (4). The last three columns are the number of lagged first-differences in the test, the results of the Breusch-Godfrey test for serial correlation, and the number of included observations.

Testing rejects cointegration under the MTAR model in Panel A and Panel B. Although the t-max test under the MTAR specification rejects the null of no cointegration at a greater than ten percent level of significance in Panel A, the Φ test fails to reject the same null (t-max test statistic of -1.84 versus a 10% critical value of -1.77, while the Φ test statistic of 5.266 is compared to the 10% critical value of 5.36). The Φ statistic, as noted by Enders and Siklos, has substantially more power than the t-max test. Therefore the results reject the MTAR specification using the Johansen error-correction term. Further, the DOLS MTAR results reject convergence, given that the parameter ρ_2 is positive. Therefore, no further testing is required in MTAR specification, since results do not indicate cointegration.

Table 2: Asymmetric Cointegration Tests

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Johansen Error-Correction Term							
	ρ_1	ρ_2	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$	k	nR^2 (p=4)	n
TAR	**-0.089**	-0.091	*5.093*	0.001	1	2.109	228
	-2.16	-2.41		[0.981]		[0.716]	
MTAR	-0.105	-0.074	5.266	0.331	1	1.874	228
	-2.72	*-1.84*				[0.759]	
Panel B: Dynamic OLS Error-Correction Term							
	ρ_1	ρ_2	$\rho_1 = \rho_2 = 0$	$\rho_1 = \rho_2$	k	nR^2 (p=4)	n
TAR	-0.165	**-0.056**	*5.210*	2.295	1	1.13	231
	-2.46	-2.14		[0.130]		[0.889]	
MTAR	-0.189	0.081	—	—	2	2.042	230
	-6.10	2.29				[.728]	

In all tests, the attractor, τ, is set equal to zero. Columns (1) and (2) give the coefficient estimates from equations (XX+1) and (XX+2), and the t-statistics are provided. Column (5) is the number of lagged first-difference terms included as specified by the Schwarz criterion and serial correclation tests. Column (6) is the critical value under the Breusch-Godfrey serial correclation test and its associated p-value is contained in the brackets below the critical value. Column (7) is the included number of observations. Bold indicates rejction at a 5% significance level and italics indicated rejection at a 10% significance level.

Results in Table 3, however, generally support cointegration between consumption, labor income, and wealth in the presence of asymmetry. The t-max test suggests convergence and cointegration at a five percent level of significance when testing either error-correction term under the TAR model, and the Johansen error-correction term under the M-TAR model. The Φ test also rejects the null of no cointegration, but at smaller level of confidence (5.093 test statistic for the Johansen error-correction term and 5.210 test statistic for the DOLS error-correction term, compared to 4.92 critical value at the 10% significance levelInterestingly, the testing fails to reject symmetry in the error-correction term. In column (4), the equality of ρ_1 and ρ_2 is supported, although the p-value using the DOLS error-correction term (0.13) is marginal.

Vector Error Correction Models

Two vector error correction models (VECM) are specified, given the TAR results. Both VECMs use the threshold of τ equal to zero to examine the potential asymmetric adjustment of consumption, labor income, and wealth. In equation (4), x represents a (3 X 1) vector of variables, α is a (3 X 1) vector of adjustment parameter, and β is a (1 X 3) vector of long run (cointegrating) parameters derived in Table 1. The construction of the Heaviside indicator, I, is also indicated in equation (4), and Δ is the first-difference operator. Finally, γ is a (3 X 3) matrix of coefficients, and ω is a (3 X 1) vector of constants. The results from equation (4) are presented in Table 3. Panel A presents the results using the Johansen error correction term, and Panel B provides results using the DOLS error correction term.

(4)

The results in Panels A and B generally support those of Lettau and Ludvigson (2004). The value of the consumption and labor income adjustment parameters, when controlling for asymmetric adjustment, are economically small and not statistically different from zero. Further, tests fail to reject the null of joint statistical insignificance of the consumption and labor income adjustment parameters. Therefore, consumption and labor income are weakly exogenous in the results presented in Table 3 (in the appendix). Additionally, both sets of results narrowly fail to reject the null of no short run Granger Causality of wealth on consumption, since the sum of the parameters on the lagged first-difference terms of wealth are statistically zero.

In contrast to the consumption and labor income results, the wealth adjustment parameters are economically large and statistically significant. Tests also reject the joint insignificance of the adjustment parameters. Therefore, wealth adjusts to restore long-run equilibrium in response to a shock that disrupts equilibrium. Concerning the issue of asymmetry, tests fail to reject the equality of the two wealth adjustment parameters. Therefore, the long run adjustment of wealth appears to be symmetric around equilibrium.

Conclusion

The traditional set of tests for cointegration fail to reject the null of no cointegration between consumption, labor income, and wealth. These tests are known to suffer from low power in the presence of asymmetric adjustments. Tests do, however, support cointegration when allowing for asymmetric long run adjustment. Contrary to previous literature, when estimating the potential asymmetries in a vector error correction model, results fail to reject symmetry in the adjustment of wealth. Further, tests fail to reject weak exogeneity of consumption and labor income. Therefore, wealth appears to adjust to shocks in the long run. Therefore, results here generally support those of Lettau and Ludvigson (2004). Consumption, in the specification presented here, does not exhibit asymmetric adjustment to shocks in the long run.

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Notes

1. See Lettau and Ludvigson for the caveats associated with their data. Notably, household net worth includes financial wealth, housing wealth, and consumer durables, as durables are assumed to represent investments in capital stock.
2. The date of 1981 was determined by a Zivot and Andrews (1992) test for structural break.
3. Taxes are calculated as [wages and salaries/(wages and salaries + Proprietors' income with inventory valuation and capital consumption adjustments + Rental income of persons with capital consumption adjustment + Personal income receipts on assets)]*Personal current taxes.
4. Using the Engle-Granger methodology, results also fail to reject the null of no cointegration.

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