Management Science

Publication details, including instructions for authors and subscription information:
http://pubsonline.informs.org

Changes in the Composition of Publicly Traded Firms: Implications for the Dividend-Price Ratio and Return Predictability

Stephan Jank

To cite this article:
Stephan Jank (2015) Changes in the Composition of Publicly Traded Firms: Implications for the Dividend-Price Ratio and Return Predictability. Management Science 61(6):1362-1377. http://dx.doi.org/10.1287/mnsc.2013.1883

MANAGEMENT SCIENCE
Vol. 61, No. 6, June 2015, pp. 1362–1377
ISSN 0025-1909 (print) | ISSN 1526-5501 (online)

http://dx.doi.org/10.1287/mnsc.2013.1883

Changes in the Composition of Publicly Traded Firms: Implications for the Dividend-Price Ratio and Return Predictability

Stephan Jank
Finance Department, Frankfurt School of Finance and Management, 60314 Frankfurt am Main, Germany; and
Centre for Financial Research (CFR), 50923 Cologne, Germany, s.jank@fs.de

This paper documents how the changing composition of U.S. publicly traded firms has prompted a decline in the long-run mean of the aggregate dividend-price ratio, most notably since the 1970s. Adjusting the dividend-price ratio for such changes resolves several issues with respect to the predictability of stock market returns: the adjusted dividend-price ratio is less persistent, in-sample evidence for predictability is more pronounced, there is greater parameter stability in the predictive regression (particularly during the 1990s), and there is evidence of out-of-sample predictability.

Data, as supplemental material, are available at http://dx.doi.org/10.1287/mnsc.2013.1883.

Keywords: return predictability; dividend-price ratio; sample selection
History: Received March 26, 2013; accepted October 30, 2013, by Itay Goldstein, finance. Published online in *Articles in Advance* May 27, 2014.

1. Introduction

The dividend-price (d-p) ratio has a long tradition as a predictive variable for market returns (e.g., Campbell and Shiller 1988, Fama and French 1988, Campbell 1991, Hodrick 1992), though in recent work several issues emerged regarding its use in predictive regressions. First, the d-p ratio exhibits strong persistence, which poses statistical problems in the predictive regression (Nelson and Kim 1993, Stambaugh 1999, Ferson et al. 2003, Valkanov 2003, Ang and Bekaert 2007). Second, the parameters of predictive regressions are unstable over time (Viceira 1997, Paye and Timmermann 2006), and accordingly, the in-sample predictability of returns seemed to disappear around the mid-1990s. Third, the out-of-sample performance of the return forecasting regression is rather poor (Bossaerts and Hillion 1999; Goyal and Welch 2003, 2008).[1] Lettau and Van Nieuwerburgh (2008) reconcile these observations with return predictability by allowing for steady-state shifts in the d-p ratio. They provide strong evidence for structural breaks in the long-run mean of the d-p ratio and cite various possible reasons for these shifts, including persistent improvements in risk sharing, changes in the long-run growth rate of the economy, or changes in tax code or payout policy.

Building on Lettau and Van Nieuwerburgh's (2008) work, this paper explores a novel channel for steady-state shifts in the d-p ratio, namely, systematic changes in the composition of publicly listed corporations. I start by comparing the d-p ratio of all domestic corporations, including both publicly traded and closely held firms, against that of publicly traded firms. Over time, the all-domestic-equity d-p ratio increasingly differs from that of publicly traded firms. This divergence can be explained by changes in the composition of publicly listed firms. Systematic differences in the dividends paid by entering and exiting firms have prompted a decline in the steady state of the observed d-p ratio, particularly since the 1970s and 1980s. It appears that as it became easier for firms to go public, such as through NASDAQ, this route came to be used overproportionately by low-dividend-paying firms in need of capital. The decline since the 1980s also may be associated with the growing importance of S corporations as an organizational form, which provided tax incentives for high-dividend-paying firms to be privately held, thereby leading to a population of public firms with lower average d-p ratios. To account for these composition changes, I subtract the cumulative steady-state change as a result of entering or exiting firms from the ordinary d-p ratio of publicly traded firms.

In the second part of this paper, I compare the all-domestic-equity d-p ratio and the d-p ratio adjusted for composition changes with the commonly used d-p ratio with respect to the aforementioned issues.

[1] See Lettau and Ludvigson (2010) and Koijen and Van Nieuwerburgh (2011) for recent surveys.

The results can be summarized as follows. The all-domestic-equity and composition changes-adjusted d-p ratio are more mean reverting. In-sample evidence for return predictability is stronger and also present after correcting estimates for small-sample bias. Furthermore, the forecasting relation remains stable over time, especially throughout the 1990s, and there is evidence for out-of-sample predictability. The composition changes-adjusted d-p ratio also performs well compared with the repurchase-adjusted d-p ratio proposed by Boudoukh et al. (2007). The repurchase-adjusted ratio does not add any information about future market returns, once the d-p ratio is adjusted for composition changes. Moreover, when predicting market returns in the full sample, there is no evidence for predictability applying the repurchase adjustment, but there is evidence for predictability using the adjustment for composition changes.

This paper contributes to the following fields of research. First, it relates to Lettau and Van Nieuwerburgh (2008), who provide evidence for structural breaks in the d-p ratio and show that by adjusting for these breaks, return predictability can be improved. Although this purely econometric approach can identify breaks ex post, it has difficulties in estimating the timing and, more importantly, the magnitude of a break in real time, which results in a poor out-of-sample performance. Therefore, to detect structural breaks in real time, it is essential to understand the economic mechanism behind these shifts. Changes in the composition of publicly traded firms provide an economic explanation for the observed shifts in the steady state of the d-p ratio. The adjustment for composition changes can be calculated in real time, and thus investors can exploit the predictability relationship (Goyal and Welch 2003, 2008).

A large body of literature studies the decline of dividends since the 1970s (e.g., Fama and French 2001, DeAngelo et al. 2004, Leary and Michaely 2011). Two broad trends can be observed: on the one hand, payout policy has changed, as corporations increasingly make use of repurchases; on the other hand, there is a structural change in the population of listed firms mostly driven by newly listed firms with a low propensity to pay dividends. Repurchases and composition changes are largely independent phenomena as repurchases are predominantly used by firms that already pay dividends (Fama and French 2001).

Boudoukh et al. (2007) investigate one of these trends, the growing importance of repurchases, regarding its effect on the d-p ratio and return predictability. They argue that dividends alone might not fully capture cash flows to investors and thus amend the d-p ratio for repurchases, which improves evidence of return predictability. However, whether a repurchase-adjusted d-p ratio is appropriate is open to debate in the literature. For example, Cochrane (2008) argues that there is nothing wrong with the d-p ratio in an accounting sense, as repurchases merely delay the eventual payment of dividends. Koijen and Van Nieuwerburgh (2011) reason that whether a repurchase adjustment is appropriate depends on whether or not one considers an investor who participates in every stock repurchase. In the end, it remains an empirical question whether changes in payout policy are relevant for asset prices (Boudoukh et al. 2007). Using an extended sample (1926–2009), Koijen and Van Nieuwerburgh (2011) find no evidence for return predictability despite the repurchase adjustment when cash-reinvested dividends (Chen 2009, Van Binsbergen and Koijen 2010) instead of market-reinvested dividends are used.

Whereas various modifications in payout measures have been extensively studied (other studies include, e.g., Bansal and Yaron 2007 and Larrain and Yogo 2008), the effect of composition changes on the d-p ratio and return predictability has, to the best of my knowledge, not been addressed yet. The current study fills this gap by providing a comprehensive analysis of how systematic differences in entering and exiting firms have led to a decline in the steady state of the d-p ratio. I show that when composition changes are taken into account, the common d-p ratio predicts market returns as implied by the present value identity.

2. Steady-State Shifts in the Dividend-Price Ratio of Publicly Traded Corporations

The theoretical motivation for predictive regressions is the log-linear approximation of the present value relationship by Campbell and Shiller (1988), which Lettau and Van Nieuwerburgh (2008) extend to allow for time-varying steady-state growth rates of dividends and returns:

$$dp_t = \overline{dp}_t + \mathbb{E}_t \sum_{j=1}^{\infty} \rho_t^{j-1} \big[(r_{t+j} - \bar{r}_t) - (\Delta d_{t+j} - \overline{\Delta d}_t) \big], \quad (1)$$

where dp_t is the dividend-price ratio, r_t is the market return, and Δd_t refers to dividend growth. All lowercase letters denote variables in logs. The expectation operator conditional on information at time t is denoted by $\mathbb{E}_t$, and $\rho_t = 1/(1+\exp(\overline{dp}_t))$. The long-term means of the d-p ratio, return, and dividend growth are denoted by an overscore: $\overline{dp}_t$, $\bar{r}_t$, and $\overline{\Delta d}_t$, respectively, where the time index indicates that the steady state can change over time. Equation (1) states that deviations of the d-p ratio from its steady state should forecast either future returns or dividend growth, or both. This is generally the motivation for the predictive regressions of returns and dividend growth:

$$r_{t+1} - \bar{r}_t = \beta^r (dp_t - \overline{dp}_t) + \varepsilon^r_{t+1} \quad \text{and} \tag{2}$$

$$\Delta d_{t+1} - \overline{\Delta d}_t = \beta^d (dp_t - \overline{dp}_t) + \varepsilon^d_{t+1}, \tag{3}$$

though this study focuses on Equation (2), the return predictability equation.

Lettau and Van Nieuwerburgh (2008) show that changes in the long-run mean of the d-p ratio can explain the high persistence of the d-p ratio; if not taken into account, these shifts distort the predictive regression, resulting in parameter instability and poor out-of-sample predictability. The authors find strong evidence of structural breaks in the d-p ratio, so they suggest using regime-specific means to demean the d-p ratio. Although structural breaks in the d-p ratio can be identified, the economic explanation for these changes is still unresolved. Lettau and Van Nieuwerburgh (2008) cite several possible explanations: improvements in risk sharing, changes in the long-run growth rate of the economy (e.g., Lustig and Van Nieuwerburgh 2005, Krueger and Perri 2006, Lettau et al. 2008), changes in tax code (e.g., McGrattan and Prescott 2005), or changes in payout policy (e.g., Fama and French 2001, Grullon and Michaely 2002, Brav et al. 2005, Boudoukh et al. 2007).

I explore a new type of channel for steady-state shifts in the d-p ratio. Shifts can be caused by composition changes in the firms, which the researcher observes. When estimating predictive regressions as in Equations (2) and (3), the d-p ratio and returns are usually measured with a broad stock market index, such as the S&P Composite or all stocks traded on the New York Stock Exchange (NYSE), American Stock Exchange (AMEX), NASDAQ, and Arca—that is, the universe of the Center for Research in Security Prices (CRSP) stock database. The sample is thus usually restricted to publicly traded firms, whereas privately held firms remain unobservable. In this setting, observed and unobserved firms may differ systematically in the amount of dividends they pay. This difference between observed and unobserved firms on its own is not a problem as long as there are no changes in composition over time. If the composition of observed firms changes over time, though, the long-run mean of the observed d-p ratio changes while the overall d-p ratio stays constant. For example, if more firms that pay low dividends go public, which makes them observable to the researcher, the *observed* d-p ratio decreases, yet the *overall* d-p ratio remains stationary around its steady state, ceteris paribus. In this study I explore this potential channel for nonstationary shifts in the d-p ratio by contrasting the d-p ratio of publicly traded firms (i.e., the CRSP sample) with the overall d-p ratio of all domestic corporations.

3. Data

The annual d-p ratio of all domestic corporations is calculated by dividing all corporate dividends paid in the economy in that year by the end-of-year total market capitalization of all domestic corporations. The source of dividend data is the national income and product accounts (NIPA), which obtains the original data from the corporate income tax returns gathered by the Internal Revenue Service (IRS). I adjust this series for capital gain distributions and interest payments from regulated investment trusts (i.e., mutual funds) using the NIPA adjustment factors. The resulting figure reflects the dividends paid by all domestic corporations. (For further details on NIPA dividends, see Petrick 2002). I divide this figure by the total market capitalization of all domestic corporations including both publicly traded and privately held firms, gathered from the flow of funds accounts (FFA). The earliest availability of annual total market capitalization from the FFA is 1945, and the latest divided figures (NIPA, IRS) are from 2008. Therefore, the main analysis of this paper focuses on the postwar period (1945–2008 for the d-p ratio, 1946–2009 for returns). The appendix provides further details on the construction of the all-domestic-equity d-p ratio.

Because the market value of all firms, publicly traded and closely held, in principle is unobservable, the FFA must rely on an approximation. The primary method for estimating the overall market capitalization is the perpetual inventory approach.[2] That is, data on net equity issuances of all corporations and capital gains from a broad market index (e.g., the Dow Jones U.S. Total Market Index) are used to extrapolate the current level of total market capitalization from the previous level. This approximation admittedly has its limitations. However, when measuring the d-p ratio, there is a trade-off between two nonperfect alternatives. One possibility is to rely on traded stocks, which allows an exact measurement of market value and dividends but restricts the researcher to a specified subsample, which is subject to the aforementioned problems. The other option is to consider all corporations in the economy. In this case there is a good measure for aggregate dividends originating from tax records; however, the total market value must be approximated. I validate the quality of the FFA's approximation by comparing the relation of total market capitalization of publicly traded and all domestic firms to dividends and earnings. There is no divergence of the earnings-price ratios, which suggests that the FFA's approximation of the overall market value works well. For a detailed discussion, see §6.2.

[2] In 2008, the last year of the sample, the FFA changed its method to estimate the total market capitalization. The new method matches privately held firms with publicly traded based on industry and revenue profile to estimate the market value of privately held firms.

In the following analysis, I compare the overall d-p ratio with the d-p ratio of publicly traded firms (CRSP). Taking into account recent concerns noted by Chen (2009), Van Binsbergen and Koijen (2010), and Koijen and Van Nieuwerburgh (2011), I do not reinvest dividends at the market return but instead aggregate them over the year by summing up monthly dividends. In addition to dividend growth being less volatile (Van Binsbergen and Koijen 2010), this approach has several advantages. First, the CRSP d-p ratio is directly comparable to the all-domestic-equity d-p ratio, which sums up dividends over the year as well. Second, the market-reinvested ratio shows a higher persistence than the zero-rate-reinvested d-p ratio (Koijen and Van Nieuwerburgh 2011). Third, reinvestment at the market return leads to an overstatement of return predictability in the pre-1945 sample (Koijen and Van Nieuwerburgh 2011). The appendix provides details on the construction of the CRSP d-p ratio.

Figure 1 compares the all-domestic-equity d-p ratio with the d-p ratio of publicly traded firms (CRSP). The two ratios are very similar in the first part of the sample; the all-domestic-equity d-p ratio is always slightly higher than the CRSP d-p ratio, which indicates that publicly traded firms have higher d-p ratios on average. Starting around the 1970s, the two ratios begin to diverge, and the wedge between them is increasing. More details appear in the left-hand graph of Figure 2, which plots the difference between the two ratios over time. Whereas the difference before 1970 averaged approximately −0.24, it widened to −0.31 in the 1970s, to −0.42 in the 1980s, to −0.66 in the 1990s, and to −0.87 in the years 2000–2008. In the following, I explore the degree to which new listings and delistings can account for this divergence.

4. Changes in the Composition of Publicly Traded Firms and Their Effect on the Dividend-Price Ratio

4.1. The Time-Varying Steady State of the Dividend-Price Ratio

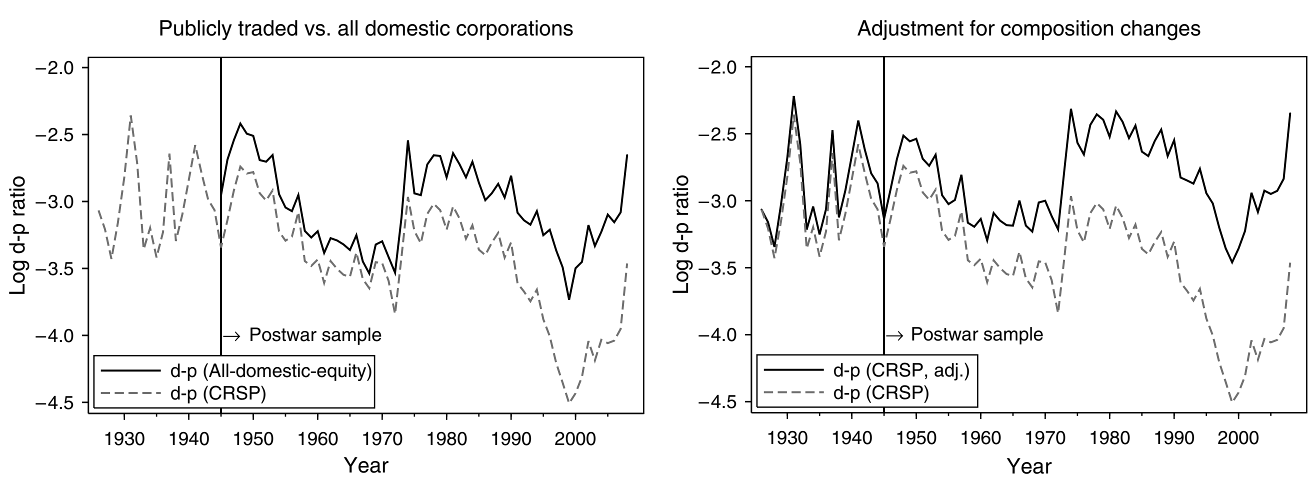
My analysis builds on work by Fama and French (2001), who find a decline in the number of dividend-paying firms starting from the 1970s. Whereas 66.5% of all publicly listed firms paid dividends in 1968, only 20.8% paid dividends in 1999. Fama and French (2001) cite two reasons: (1) an increase in the rate of dividend-paying firms that delist and, more importantly, (2) a strongly increasing share of newly listed firms that do not pay dividends. In their analysis, Fama and French (2001) focus on individual firms and, when aggregating use equal weighting, whereas in this section, I investigate the value-weighted effect of composition changes on the aggregate d-p ratio.

As a first step, I decompose the steady state of the d-p ratio into the steady state of the firms that were continuously listed from the previous to the present year $\overline{dp}_t^s$ and the change in steady state as a result of entering or exiting firms $\Delta\overline{dp}_t^e$:

$$\overline{dp}_t = \overline{dp}_t^s + \Delta\overline{dp}_t^e. \tag{4}$$

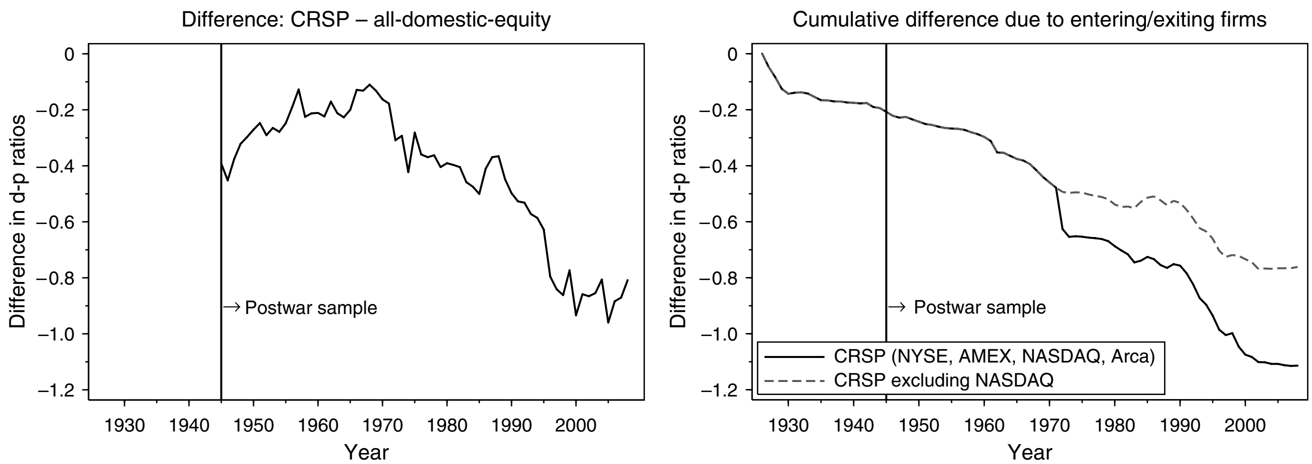
If there is no systematic difference between entering or exiting firms and the continuously listed firms with respect to their d-p ratio, $\Delta\overline{dp}_t^e$ equals zero, and all else being equal, the steady state of the d-p ratio is constant over time. Conversely, if entering or exiting firms have systematically lower or higher d-p ratios, the steady state $\overline{dp}_t$ shifts. Because the steady state of the continuously listed firms at time t equals the

Figure 1 **Comparison of Dividend-Price Ratios**



Notes. The left-hand graph shows the log dividend-price (d-p) ratio of all publicly traded corporations, i.e., traded on NYSE, AMEX, NASDAQ, or Arca (CRSP), and the d-p ratio all domestic corporations. The right-hand graph displays the CRSP d-p ratio and the d-p ratio adjusted for composition changes. The sample period is 1926–2008 at annual frequency. Data for the all-domestic-equity d-p ratio are available from 1945 onward.

Figure 2 Differences in Dividend-Price Ratios and Adjustment for Composition Changes



Notes. The left-hand graph displays the difference between the all-domestic-equity d-p ratio and the CRSP d-p ratio. The right-hand graph displays the cumulative difference in d-p ratio that is due to entering or exiting firms for the entire CRSP sample (NYSE, AMEX, NASDAQ, and Arca) and when NASDAQ firms are excluded. The sample period is 1926–2008 at annual frequency. Data for the all-domestic-equity d-p ratio are available from 1945 onward.

steady state of the previous period ($\overline{dp}^s_t = \overline{dp}_{t-1}$), we can iterate forward Equation (4) and rewrite it as

$$\overline{dp}_t = \overline{dp}_0 + \sum_{i=1}^{t} \Delta \overline{dp}^e_i. \tag{5}$$

The d-p steady state at time t is the initial steady state at time zero plus the sum of changes in steady state as a result of entering or exiting firms.

To estimate the change in the d-p steady state induced by entering or exiting firms $\Delta \overline{dp}^e_t$, I proceed as follows. I only consider firms that are continuously listed in the previous and current year and calculate their d-p ratio: dp^s_t. The difference between the d-p ratio of continuously listed firms (dp^s_t) and the d-p ratio of all firms listed in the current year (dp_t) can be attributed to composition changes ($\Delta \overline{dp}^e_t$).[3]

The right-hand graph of Figure 2 plots the cumulative sum of changes in the d-p ratio as a result of entering or exiting firms, which is the path of the time-varying steady state $\overline{dp}_t$ relative to its initial value $\overline{dp}_0$, as stated in Equation (5). With the exception of the first years, the steady state of the d-p ratio is only slightly decreasing and remains quite stable in the first part of the sample. In the latter part of the sample, however, there is a strong decline around the 1970s and in particular from the 1980s onward.

4.2. Discussion: Economic Reasons for Composition Changes in Publicly Traded Corporations

What explains systematic changes in the composition of publicly traded corporations? A possible explanation is that from the 1970s on, it became easier for firms to go public, such as through NASDAQ, which at that time was overproportionately used by low-dividend-paying firms in need of capital. To demonstrate to what degree the decline in steady state can be ascribed to the introduction of NASDAQ, I rerun the analysis, excluding all firms listed on NASDAQ. When NASDAQ data are included in the CRSP database in 1972, a strong decline occurred in the d-p steady state (see the right-hand graph of Figure 2). Furthermore, the overall CRSP sample's d-p steady state declines at a faster rate than that of the sample without NASDAQ, indicating that firms with systematically lower d-p ratios preferably chose NASDAQ when going public. However, as Figure 2 demonstrates, the overall decline in the d-p steady state cannot be explained by NASDAQ alone. These findings are consistent with prior literature. Fama and French (2001) also document that the lower number of dividend-paying firms is not restricted to NASDAQ but also is the case for AMEX and NYSE. As part of their analysis, Lettau and Van Nieuwerburgh (2008) and Koijen and Van Nieuwerburgh (2011) exclude NASDAQ firms from the sample, which improves the

[3] Another possibility to test for composition changes is to restrict the sample to those firms that survived for the entire sample period and compare this d-p ratio with the conventional d-p ratio. The number of firms that existed in 1945 and were continuously listed until 2008 (using the same unique CRSP identifier PERMNO) is 101. The d-p ratio of surviving firms is very similar to the overall CRSP d-p ratio before the 1970s (average difference of 0.02). From the 1970s onward, the ratios increasingly deviate from each with a maximum difference of −0.35. The continuously listed d-p ratio is higher than the overall CRSP d-p ratio, which suggests that because of the entry of lower-dividend-paying firms, the mean of the d-p ratio changed. However, this approach suffers from survivorship bias. Only those firms that survived or did not choose to delist are considered. Fama and French (2001) find that dividend payers delist at a higher rate in the period 1978–1999 than previously, which explains why the wedge between the series is not as pronounced as for the all-domestic-equity versus CRSP series.

return predictability results slightly, but the structural break problem remains, as Lettau and Van Nieuwerburgh (2008) note. Thus, the composition changes of publicly traded corporations were partly driven by the introduction of NASDAQ, but not entirely.

Furthermore, if the incentive to be public or private is systematically different for high- and low-dividend-paying firms, this also can result in a shift in the d-p steady state of listed firms. Beginning in the mid-1980s, tax advantages from a single level of taxation for S corporations provided an incentive for high-dividend-paying firms to be privately held, which offers a possible explanation for the large number of firms paying low dividends that became public. For "ordinary" corporations, so-called C corporations, income is first taxed at the business level, and then dividends received by shareholders are again taxed at the individual level. In contrast, S corporations are not subject to taxes at the business level, making this form of corporation attractive, especially for firms that intend to pay large dividends to their shareholders. Indeed, payout ratios for S corporations are generally higher than those of C corporations, averaging around 83.9%, compared with 55.5% for C corporations.[4] Although the single level of taxation is an attractive feature of S corporations, they also suffer additional restrictions; in particular, S corporations are restricted in their number and type of shareholders. The restrictions on who may own an S corporation effectively prevent the corporation from being publicly traded, which creates different incentives for high- and low-dividend paying firms to be publicly listed or privately held.

Tax reforms in 1986 and 1996 made S corporations more attractive, and over time, an increasing number of companies have responded to this tax incentive and have chosen to file their taxes under Subchapter S of the Internal Revenue Code (e.g., Petska 1996, Legel et al. 2003). The number of S corporations was 0.5 million in 1980 but steadily increased, such that by 1997, S corporations became the most prevalent corporation type, with 4.1 million of them operating by 2008.[5] Moreover, these corporations are not necessarily small. Increasingly, large corporations prefer to file their taxes as S corporations (McKinnon 2010). The rising number of S corporations is also important in economic terms. Before 1986, the amount of earnings generated by S corporations was small, averaging approximately 2.7% of all corporate earnings. However, it sharply increased after the tax reform of 1986 to 10.7% and steadily rose in the following decades to 18.5% in the 1990s and 30.8% in 2000–2008. The share of dividends paid by S corporations (available only since 1991) shows a similar pattern: S corporations generated around 17.7% of all dividends in 1991, and the percentage rose to 42.9% in 2008. In summary, the single level of taxation made S corporations increasingly popular from the mid-1980s onward. The tax advantage of S corporations creates different incentives for high- and low-dividend-paying firms to be privately held or publicly traded, which provides a possible explanation for the large share of low-dividend-paying firms becoming public from the mid-1980s onward.

Other factors may play a role in the decision of whether to be listed. For example, changes in the regulatory framework such as disclosure requirements could encourage or discourage public listing, possibly having a systematic effect on the composition of publicly traded firms. It is beyond the scope of this paper, however, to explore all determinants of composition changes. Moreover, in the future there might be other reasons that lead to systematic composition changes. The adjustment method for the dividend-price ratio proposed in the next section can be applied independently.

4.3. Adjusting the Dividend-Price Ratio for Composition Changes

When looking at Figure 2, we see that the cumulative change in the d-p ratio as a result of entering or exiting firms (right-hand graph) matches the difference between the publicly traded and the overall d-p ratio (left-hand graph) both in magnitude and in pattern. The overall change in the mean of the d-p ratio is also similar to the magnitude of the structural break reported by Lettau and Van Nieuwerburgh (2008), who estimate a break of −0.86 in 1991.

Now that we have estimated the time-varying steady state as a result of entering or exiting firms $\overline{dp}_t$, we can appropriately demean the d-p ratio as required by the present value formula: $dp_t - \overline{dp}_t$. The mechanism of the adjustment is in spirit similar to Lettau and Van Nieuwerburgh (2008), but instead of a regime switch, it is a continuous adjustment. The resulting d-p ratio (adjusted for composition changes), along with the unadjusted CRSP d-p ratio, is plotted in the right-hand graph of Figure 1, and the series is quite similar to the all-domestic-equity d-p ratio plotted in the left-hand graph. As is apparent from the graph, the adjusted series is less persistent than the unadjusted one.

The d-p adjusted for composition changes is not an extension of the present value relationship; rather, it constitutes a better measurement of the d-p ratio's temporary deviations from its steady state. At each point in time, the Campbell and Shiller (1988) approximation of the present value formula holds for the exact portfolio of all publicly traded corporations. A temporary deviation of the d-p ratio from its steady state at time t should forecast either future returns or dividend

[4] Figures are averages over the period 1991–2008 based on IRS statistics.

[5] *Source.* IRS, Statistics of Income Tax Stats—Integrated Business Data (http://www.irs.gov/uac/SOI-Tax-Stats-Integrated-Business-Data).

growth, or both. However, when solely looking at the d-p ratio of publicly traded firms, a decline in the d-p ratio from time t to $t+1$ can mean two things: it is either a transitory divergence or a permanent change in the steady state. The latter holds if low-dividend-paying firms go public or high-dividend-paying firms go private. Looking only at the simple d-p ratio, the decline as a result of composition changes is mistakenly attributed to a temporary divergence of the d-p ratio from its steady state, even though it is actually a persistent change of the d-p steady state. The procedure described above captures the persistent change of the steady state and thus provides a way to appropriately demean the d-p ratio.

Even though the all-domestic-equity d-p ratio is useful when the goal is to understand the underlying economic process of structural shifts, it is disadvantageous in that macroeconomic data—in particular, tax data from the IRS—are only published after a delay. In contrast, the d-p ratio adjusted for composition changes can be computed without a time lag. Thus, an out-of-sample forecast in real time is not feasible for the all-domestic-equity d-p ratio but is implementable for the adjusted d-p series.

Next, I compare the time-series properties and ability to predict future market return of the CRSP d-p ratio, the all-domestic-equity d-p ratio, and the CRSP d-p ratio adjusted for composition changes. The main focus is on the post-1945 sample, for which all three d-p ratios are available. The sample extended to 1926, for which only the CRSP d-p ratio and the CRSP d-p ratio adjusted for composition changes are available, serves as a robustness check.

5. Comparison of Different Dividend-Price Ratios

5.1. Descriptive Statistics

Table 1 reports the summary statistics for the three d-p ratios. The observations from Figure 1 regarding the ratios' persistence are confirmed. In the post-1945 sample, the augmented Dickey and Fuller (1979) (ADF) test fails to reject the null hypothesis of nonstationarity for the CRSP d-p ratio with a p-value of 0.35. Even though the Dickey–Fuller test cannot reject the null hypothesis of nonstationarity for the all-domestic-equity and adjusted d-p ratio in the postwar period, p-values are considerably lower. Constraining the analysis to a smaller subsample naturally reduces the test's power to reject the null. In the full sample from 1926 to 2008, the Dickey–Fuller test yields a result comparable to that of Lettau and Van Nieuwerburgh (2008). The null hypothesis of nonstationarity is not rejected for the unadjusted CRSP d-p ratio but is rejected for the adjusted one with a p-value of 0.010.

Table 1 Summary Statistics

Ratio	Mean	SD	AC(1)	AC(2)	ADF	p-value
	Sample period: 1945–2008					
d-p (CRSP)	−3.48	0.43	0.91	0.82	−1.86	(0.351)
d-p (all-domestic-equity)	−3.04	0.31	0.82	0.70	−2.17	(0.216)
d-p (CRSP, adjusted)	−2.85	0.30	0.80	0.66	−2.16	(0.220)
	Sample period: 1926–2008					
d-p (CRSP)	−3.37	0.44	0.87	0.75	−2.42	(0.136)
d-p (CRSP, adjusted)	−2.86	0.30	0.70	0.47	−3.43	(0.010)

Notes. Provided are descriptive statistics for the three different dividend-price ratios: d-p (CRSP) is the dividend-price ratio of all corporations traded on NYSE, AMEX, NASDAQ, and Arca based on the CRSP database; d-p (all-domestic-equity) covers all corporations, i.e., publicly traded and privately held, available from 1945; and d-p (CRSP, adjusted) is the d-p ratio adjusted for composition changes in publicly traded corporations. For each variable, the mean, the standard deviation, first- and second-order autocorrelations (AC(1) and AC(2)), and the augmented Dickey–Fuller (ADF) test and its p-values are reported. The sample period is 1926–2008 at annual frequency.

The reduction in persistence is also evident when we compare the autocorrelation of the series. Whereas the first-order autocorrelation of the commonly used d-p ratio is 0.91, it is 0.82 and 0.80 for the all-domestic-equity and the adjusted series in the postwar sample, respectively. In the overall sample, the first-order autocorrelation is further reduced to 0.87 (CRSP) and 0.70 (CRSP, adjusted for composition changes). Note that the autocorrelation of the all-domestic-equity d-p ratio and the CRSP d-p ratio adjusted for composition changes is considerably lower than those commonly used in the literature, so the time series is much better behaved. For example, Ferson et al. (2003) show by means of simulation that for regressors with autocorrelation coefficients less than or equal to 0.90, no serious spurious regression bias arises in the t-statistics or in the R^2. Comparable to the break-adjusted series of Lettau and Van Nieuwerburgh (2008), I find that the standard deviation of the all-domestic-equity and adjusted d-p ratio is lower than that of the unadjusted CRSP series, with a reduction of about one-third.

5.2. In-Sample Predictability

In the following, I compare the different d-p ratios with respect to their ability to predict market returns. I start with the in-sample predictive regression over the postwar sample, ranging from 1945 to 2008 for d-p ratios and from 1946 to 2009 for market returns; then in a second step, I continue to analyze the extended sample ranging from 1926 to 2008 for d-p ratios and 1927 to 2009 for market returns. Table 2 displays the ordinary least squares (OLS) regression results of market return, market excess return, and dividend growth on different lagged d-p ratios, all in logs. The table provides the slope coefficient β and the p-value for the null hypothesis $\beta^r = 0$ against the alternative $\beta^r > 0$ for the return predictability regression and $\beta^d = 0$ versus

Table 2 Predictive Regressions

Panel A: Return predictability									
	β^r	p-value (NW)	R^2	β^r_c	p-value (BS)	R^2_c	R^2 p-value (BS)	ϕ_c	ρ
Sample period: 1945–2009									
Return$_{t+1}$									
d-p$_t$ (CRSP)	0.13	(0.007)	10.76	0.08	[0.140]	8.89	[0.034]	0.97	−0.92
d-p$_t$ (all-domestic-equity)	0.21	(<0.001)	14.24	0.16	[0.045]	13.16	[0.005]	0.90	−0.90
d-p$_t$ (CRSP, adjusted)	0.21	(<0.001)	13.60	0.16	[0.050]	12.52	[0.007]	0.90	−0.92
Excess return$_{t+1}$									
d-p$_t$ (CRSP)	0.13	(0.009)	9.55	0.08	[0.155]	7.68	[0.052]	0.97	−0.90
d-p$_t$ (all-domestic-equity)	0.20	(<0.001)	13.07	0.16	[0.052]	11.99	[0.008]	0.90	−0.89
d-p$_t$ (CRSP, adjusted)	0.17	(0.001)	9.13	0.12	[0.091]	8.04	[0.031]	0.90	−0.91
Sample period: 1926–2009									
Return$_{t+1}$									
d-p$_t$ (CRSP)	0.08	(0.082)	3.19	0.05	[0.194]	2.43	[0.157]	0.91	−0.84
d-p$_t$ (CRSP, adjusted)	0.17	(0.003)	6.30	0.14	[0.052]	5.67	[0.030]	0.77	−0.84
Excess return$_{t+1}$									
d-p$_t$ (CRSP)	0.09	(0.053)	4.12	0.06	[0.151]	3.36	[0.104]	0.91	−0.84
d-p$_t$ (CRSP, adjusted)	0.14	(0.009)	4.61	0.11	[0.085]	3.98	[0.067]	0.77	−0.85
Panel B: Dividend growth predictability									
	β^d	p-value (NW)	R^2	β^d_c	p-value (BS)	R^2_c	R^2 p-value (BS)	ϕ_c	ρ
Sample period: 1945–2009									
Dividend growth$_{t+1}$									
d-p$_t$ (CRSP)	0.01	(0.699)	0.56	0.02	[0.873]	−0.23	[0.571]	0.97	0.33
d-p$_t$ (all-domestic-equity)	0.01	(0.610)	0.18	0.01	[0.725]	−0.58	[0.736]	0.90	0.25
d-p$_t$ (CRSP, adjusted)	−0.02	(0.274)	0.87	−0.01	[0.263]	0.12	[0.467]	0.90	0.31
Sample period: 1926–2009									
Dividend growth$_{t+1}$									
d-p$_t$ (CRSP)	−0.08	(0.064)	8.49	−0.07	[0.031]	7.90	[0.012]	0.91	0.40
d-p$_t$ (CRSP, adjusted)	−0.16	(0.011)	15.75	−0.16	[0.005]	15.19	[<0.001]	0.77	0.38

Notes. Reported are the OLS estimation results for the following predictive regression framework using different measures of the dividend-price ratio dp_t:

$$r_{t+1} = \alpha^r + \beta^r dp_t + \varepsilon^r_{t+1} \quad \Delta d_{t+1} = \alpha^d + \beta^d dp_t + \varepsilon^d_{t+1} \quad dp_{t+1} = \theta + \phi dp_t + \eta_{t+1},$$

where r_{t+1} is either log market return or log market excess return (panel A) and Δd_{t+1} log dividend growth (panel B). Here, d-p (CRSP) is the dividend-price ratio of all publicly traded corporations, d-p (all-domestic-equity) covers publicly traded and privately held corporations, and d-p (CRSP, adjusted) is the d-p ratio adjusted for composition changes in publicly traded corporations. The estimate of the slope coefficient β, the p-value (based on Newey–West (NW) standard errors using one lag) for the one-sided test of $\beta^r = 0$ versus $\beta^r > 0$ (panel A) and $\beta^d = 0$ versus $\beta^d < 0$ (panel B), and the regression R^2 in percentage terms are shown. Furthermore, this table reports β-coefficients and R^2 values adjusted for small-sample bias along with the p-values for a one-sided test obtained from bootstrapped (BS) distributions in brackets. Also shown are the bias-corrected estimate of the autoregressive parameter ϕ_c of the different d-p ratios and the correlation ρ of the innovations η_{t+1} and ε^r_{t+1} (panel A) and η_{t+1} and ε^d_{t+1} (panel B), respectively. The sample period is 1926–2009 at annual frequency.

$\beta^d < 0$ for the dividend growth predictability regression based on Newey and West (1987) standard errors.

In the post-1945 sample, the regression of returns on the CRSP d-p ratio yields a slope coefficient β^r of 0.13, which is significant at conventional significance levels, and the R^2 is 10.76%. Both the all-domestic-equity and the adjusted CRSP d-p ratio provide a considerable improvement in the predictive regression. For example, the slope coefficient increases to 0.21 and the R^2 to 14.24% when the all-domestic-equity d-p ratio is used. Furthermore, statistical significance is more pronounced. These results are in line with Lettau and Van Nieuwerburgh (2008), who find a comparable improvement in the predictive regression when they adjust for structural breaks. Consistent with prior literature (e.g., Chen 2009), I find no evidence of dividend growth predictability in the postwar sample.[6]

In the extended sample (1926–2009), evidence for return predictability is weaker with the CRSP d-p ratio, and there is evidence for dividend predictability. When I employ the d-p ratio adjusted for composition changes, return predictability is more pronounced, in both statistical and economic terms, compared with

[6] Van Binsbergen and Koijen (2010) find dividend growth predictability for the postwar period when using a latent variables approach, but not with OLS.

the unadjusted series. However, the explained return variation is still lower than in the postwar sample. Furthermore, evidence for dividend growth predictability in the extended sample is strengthened when using the d-p ratio adjusted for composition changes.

An important issue when dealing with the d-p ratio as a predictive variable is its strong persistence, which biases the slope coefficient in finite samples because innovations of the predictor variable are correlated with return and dividend growth innovations (Kendall 1954, Nelson and Kim 1993, Stambaugh 1999). Similarly, the R^2 suffers an upward bias in this setting. To address this issue, I calculate the small-sample bias correction proposed by Stambaugh (1999). The bias-corrected slope coefficients β_c^r and β_c^d, along with their p-values, can also be found in Table 2. Inference is based on a residual resampling bootstrap following Nelson and Kim (1993), Mark (1995), Kothari and Shanken (1997), and Kilian (1999). The table also reports the bias-corrected coefficient of determination R_c^2, which is obtained by subtracting the median value generated by the residual resampling bootstrap (Mark 1995), as well as its bootstrapped p-value. In addition, the table provides the estimates for the bias-adjusted autoregressive coefficient ϕ_c (Kendall 1954) of the predictor variable and the correlation ρ of predictor variable innovations with return and dividend growth innovations.

In the postwar sample, the slope coefficient of the conventional d-p ratio is considerably reduced through the bias correction by around 38%, to 0.08 for returns and excess returns. In both cases, the regressors are no longer statistically significant at the 10% significance level, which is in line with the findings of Stambaugh (1999). The bias correction also reduces the slope coefficient of the overall and adjusted d-p ratio but the reduction is less, ranging around 22%–28%. Moreover, the overall and the adjusted d-p ratio continue to be significant. The results for the extended sample (1926–2009) are qualitatively the same. The lower bias can be attributed to the lower autocorrelation ϕ_c of the all-domestic-equity and composition changes adjusted series, whereas the correlation of innovations ρ is very similar across different series. The coefficient of determination is affected in a similar manner, where the bias is larger for the unadjusted d-p ratio than for the adjusted ratio or the overall ratio. In particular, for the full sample, the R^2 is not statistically different from zero for the unadjusted series but is significant in the case of the composition changes-adjusted series. The small-sample bias is less severe for the dividend growth predictability regression because the correlation of innovations is considerably lower. In general, the estimates of the predictive regressions are similar for returns or excess returns. For this reason, I focus on returns in the rest of the analysis.

5.3. Parameter Stability of the Predictive Regression

Viceira (1997) and Paye and Timmermann (2006) document considerable instability of the forecasting relationship over time. To evaluate the stability of the forecasting relationship, I run rolling regressions for the different d-p ratios. To ensure the results are comparable to those from previous studies (Lettau and Van Nieuwerburgh 2008, Koijen and Van Nieuwerburgh 2011), I closely follow their settings, using a rolling window of 30 years, and report the time-varying slope coefficient estimates (for market returns) plus or minus one standard deviation. The results of the rolling regressions are displayed in Figure 3.

Similar to previous studies, the slope coefficient of the CRSP d-p ratio varies considerably over time, with a maximum of 0.45 and a minimum of 0.04 in the late 1990s. The all-domestic-equity d-p ratio instead shows a stable slope coefficient in the forecasting regression, fluctuating only slightly around the overall estimate of 0.21, even throughout the 1990s. A similar improvement in parameter stability can be observed when the d-p ratio is adjusted for composition changes. In particular, in the post-1945 sample period, the slope coefficient is fairly stable over time and comparable with that of the all-domestic-equity d-p ratio. As already seen in Figure 1, the adjustment for composition changes matters mostly in the postwar sample. There also is no substantial difference in the rolling regressions between the adjusted and unadjusted CRSP d-p ratios in the prewar sample. Therefore, the weaker evidence for return predictability in this period likely can be traced back to other structural changes, such as the shift from dividend to return predictability documented by Chen (2009) and Chen et al. (2012).

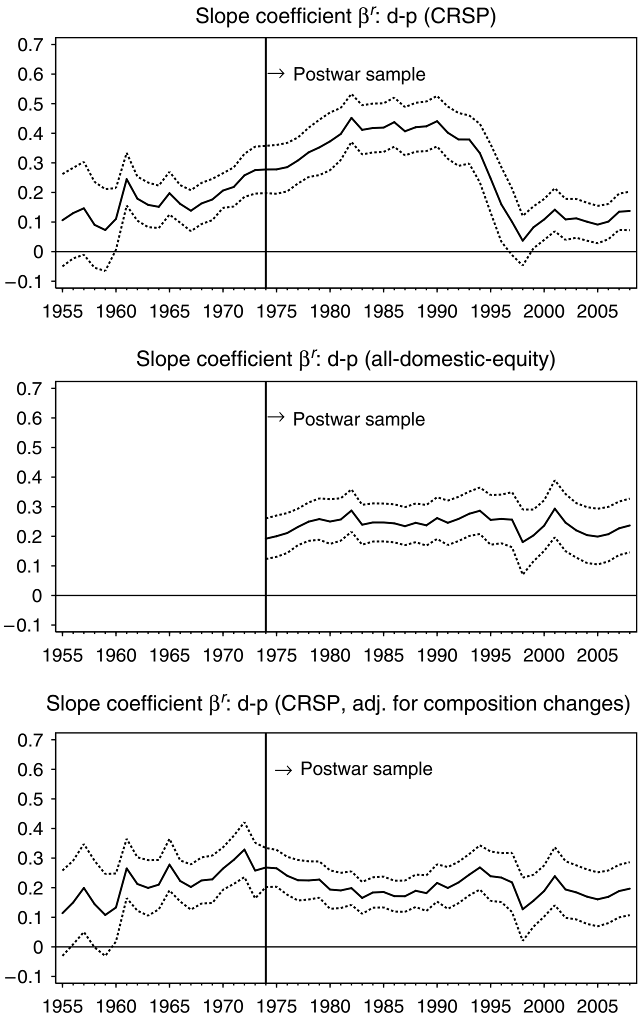
Instability in the predictive regression can be the reason for poor out-of-sample forecasts. In the following, I analyze whether and how the better parameter stability of the overall and adjusted d-p ratio manifests itself in better out-of-sample performance.

5.4. Out-of-Sample Predictability

Evidence of return predictability has recently been challenged by Goyal and Welch (2003, 2008), who find only poor out-of-sample predictability for the d-p ratio and other variables. Poor out-of-sample predictability does not necessarily contradict return predictability, because it can be attributed to the lower power of the out-of-sample tests compared with in-sample tests, as argued by Inoue and Kilian (2005) and Cochrane (2008). Furthermore, Campbell and Thompson (2008) show that prediction variables exhibit better out-of-sample predictability once restrictions are imposed on the sign of the coefficients.

Following prior literature, I compare the out-of-sample performance of the different d-p ratios to the

Figure 3 **Rolling Regression: Parameter Stability Comparison**



Notes. Shown are the results of a 30-year rolling regression of future log market return on the present log dividend-price ratio: $r_{t+1} = \alpha^r + \beta^r dp_t + \varepsilon^r_{t+1}$. The top graph shows the CRSP d-p ratio, the middle graph the all-domestic-equity d-p ratio, and the bottom graph the CRSP d-p ratio adjusted for composition changes. The graphs plot the estimate of the slope coefficient β^r (solid line) plus/minus one standard deviation (dotted lines). Standard errors are calculated by the Newey–West estimator using one lag. The sample period is 1926–2009 at annual frequency.

performance of a simple random walk model, which uses the past average market return as a naïve guess for the future market return. The initial forecasting regression contains 20 years of data, so the out-of-sample period is 1965–2009 for the postwar sample and 1946–2009 for the entire sample. I calculate the mean absolute error (MAE) and the root mean squared error (RMSE) of the predictive regression models and of the benchmark random walk model. The out-of-sample R^2 of model i is defined as $R^2_{OS,\,i} = 1 - MSE_i/MSE_{rw}$, where MSE_i is the mean squared error of model i and MSE_{rw} the mean squared error of the benchmark random walk model (see Campbell and Thompson 2008). Furthermore, I test whether the reduction in MSE of regression

Table 3 **Out-of-Sample Predictability of Return$_{t+1}$**

	MAE	RMSE	R^2_{OS}	MSE-F	p-value
Sample period: 1945–2009, out-of-sample period: 1965–2009					
Random walk	14.24	17.97	—	—	—
d-p (CRSP)	14.76	17.75	2.43	1.12	[0.080]
d-p (all-domestic-equity)	13.14	16.77	12.87	6.65	[0.004]
d-p (CRSP, adjusted)	13.28	16.95	10.99	5.56	[0.004]
Sample period: 1926–2009, out-of-sample period: 1946–2009					
Random walk	13.68	17.18	—	—	—
d-p (CRSP)	14.07	16.99	2.30	1.51	[0.050]
d-p (CRSP, adjusted)	12.95	16.31	9.94	7.07	[0.002]

Notes. Displayed are the out-of-sample predictability comparison for future market return. The out-of-sample MAE and the RMSE in percentage terms for a simple random walk model and regression models using different d-p ratios are shown. d-p (CRSP) is the dividend-price ratio of all publicly traded corporations, d-p (all-domestic-equity) covers publicly traded and privately held corporations, and d-p (CRSP, adjusted) is the d-p ratio adjusted for composition changes in publicly traded corporations. The out-of-sample R^2 of regression model i is computed as $R^2_{OS,\,i} = 1 - MSE_i/MSE_{rw}$, where MSE_{rw} is the mean squared error of the random walk model. Furthermore, this table shows the MSE-F statistic along with bootstrapped p-values, testing for equal out-of-sample MSEs of the random walk model and the respective regression model. The sample period is 1926–2009 at annual frequency.

model i compared with the random walk model is significant in statistical terms using McCracken's (2007) MSE-F statistic: $MSE\text{-}F = T \cdot (MSE_{rw} - MSE_i)/MSE_i$, where T is the size of the out-of-sample period. Inference regarding the MSE-F statistic is based on a residual resampling bootstrap. As mentioned previously, the out-of-sample results for the all-domestic-equity d-p ratio are merely a pseudo-out-of-sample forecast because of the delayed publication of macroeconomic data; the adjusted d-p ratio, in contrast, can be calculated in real time and used for real out-of-sample forecasts.

Table 3 compares the out-of-sample forecast performance of the different d-p ratios. Overall, there is only poor evidence of out-of-sample predictability when the ordinary CRSP d-p ratio is used. There is a slight reduction in RMSE compared with the random walk model, which is also significant at the 10% level; however, economically speaking, the out-of-sample R^2 is rather small.[7] Moreover, the MAE points in the other direction, favoring the naïve random walk model.

Evidence of out-of-sample return predictability is much clearer when the overall d-p ratio and the d-p ratio adjusted for composition changes are used. Both ratios show a reduction in RMSE and MAE for all

[7] Goyal and Welch (2008) and Lettau and Van Nieuwerburgh (2008) find that the CRSP d-p ratio has a higher RMSE compared with the random walk model. The slightly better performance of the unadjusted CRSP d-p ratio in Table 3 can be attributed to the reinvestment of dividends at a zero rate. Using the CRSP d-p ratio with market-reinvested dividends to predict returns yields a RMSE of 17.77 for the out-of-sample period 1946–2009, which, as in the aforementioned studies, is higher than that of the random walk model with a RMSE of 17.18.

sample periods. The forecast improvement is evident in both statistical and economic terms. For example, when predicting returns in the postwar sample, the adjusted d-p ratio reduces the RMSE of the benchmark model by −1.02 percentage points, and the p-value of the MSE-F statistic is well below the 1% significance level. This reduction in mean squared prediction error corresponds to an out-of-sample R^2 of 11%.

5.5. Long-Horizon Predictability

Much literature on return predictability, starting with Campbell and Shiller (1988) and Fama and French (1988), investigates return predictability over longer horizons finding that returns become more predictable as the horizon grows. Thus, I compare the forecasting performance of the different d-p ratios over multiple horizons in the following regression setup:

$$r_{t,\,t+H} = \alpha^r(H) + \beta^r(H)\, dp_t + \varepsilon^r_{t,\,t+H}, \qquad (6)$$

where $r_{t,\,t+H}$ is the H-year return over the time period t until $t+H$. I compute standard errors following Hodrick (1992) because Ang and Bekaert (2007) show that the properties of Hodrick standard errors are superior to Newey–West standard errors in multiple horizon regressions. Moreover, I address recent concerns expressed by Boudoukh et al. (2008), who point out that even under the null hypothesis of no return predictability, overlapping return data in conjunction with persistence in the predictive variable can lead to slope coefficients and R^2 values that rise with the forecasting horizon. For this reason I additionally provide bootstrapped p-values of coefficients and R^2 values using the residual resampling procedure described previously.[8]

Table 4 shows the return predictability regression results of Equation (6) over different horizons. Similar to previous studies, I find that for the CRSP d-p ratio, the coefficient estimates and R^2 rise with the increasing horizon. Asymptotic Hodrick standard errors suggest long-horizon predictability; however, when considering bootstrapped p-values, slope coefficients become insignificant. The R^2 values lie well in the region of what can be expected under the null hypothesis of no predictability, for example, with a p-value of 0.209 for a five-year horizon in the full sample. These results are in line with the findings by Boudoukh et al. (2008).

With regard to the all-domestic-equity d-p ratio and the d-p ratio adjusted for composition changes, evidence for return predictability for multiple horizons is notably improved, particularly in the full sample. Both the magnitude of the slope coefficient $\beta^r(H)$ and the explained variation are considerably larger throughout all horizons. For example, the five-year

[8] Another approach to addressing this problem is to calculate regression coefficients and R^2 statistics implied by a one-period vector autoregressive (VAR) model (Campbell and Shiller 1988, Kandel and Stambaugh 1989, Campbell 1991, Hodrick 1992). Calculating long-horizon statistics based on a VAR model yields comparable results to the OLS estimates.

Table 4 **Long-Horizon Predictability of Return$_{t+1}$**

	Horizon H (in years)				
	1	2	3	4	5
	Sample period: 1945–2009				
d-p (CRSP)					
β^r	0.13	0.25	0.32	0.38	0.47
p-value (Hodrick)	(0.004)	(0.006)	(0.017)	(0.027)	(0.027)
p-value (BS)	[0.140]	[0.135]	[0.166]	[0.201]	[0.199]
R^2	10.8	21.6	28.1	33.3	39.7
R^2_c	8.9	17.8	22.4	25.7	30.2
R^2 p-value (BS)	[0.034]	[0.025]	[0.033]	[0.042]	[0.039]
d-p (all-domestic-equity)					
β^r	0.21	0.39	0.50	0.60	0.75
p-value (Hodrick)	(<0.001)	(<0.001)	(0.001)	(0.003)	(0.003)
p-value (BS)	[0.045]	[0.038]	[0.046]	[0.055]	[0.045]
R^2	14.2	27.0	37.0	45.4	55.9
R^2_c	13.2	24.9	33.8	41.2	50.7
R^2 p-value (BS)	[0.005]	[0.003]	[0.002]	[0.002]	[<0.001]
d-p (CRSP, adjusted)					
β^r	0.21	0.39	0.50	0.58	0.72
p-value (Hodrick)	(<0.001)	(<0.001)	(0.001)	(0.003)	(0.002)
p-value (BS)	[0.050]	[0.040]	[0.049]	[0.065]	[0.061]
R^2	13.6	25.7	35.1	40.3	48.0
R^2_c	12.5	23.6	32.0	36.3	43.1
R^2 p-value (BS)	[0.007]	[0.004]	[0.003]	[0.004]	[0.004]
	Sample period: 1926–2009				
d-p (CRSP)					
β^r	0.08	0.19	0.24	0.29	0.34
p-value (Hodrick)	(0.061)	(0.030)	(0.043)	(0.054)	(0.064)
p-value (BS)	[0.194]	[0.145]	[0.178]	[0.198]	[0.209]
R^2	3.2	7.9	9.4	11.2	13.3
R^2_c	2.4	6.4	7.2	8.2	9.6
R^2 p-value (BS)	[0.157]	[0.102]	[0.141]	[0.165]	[0.169]
d-p (CRSP, adjusted)					
β^r	0.17	0.41	0.59	0.75	0.86
p-value (Hodrick)	(0.003)	(<0.001)	(<0.001)	(<0.001)	(<0.001)
p-value (BS)	[0.052]	[0.014]	[0.008]	[0.005]	[0.005]
R^2	6.3	17.1	26.9	35.5	41.9
R^2_c	5.7	15.9	25.2	33.5	39.5
R^2 p-value (BS)	[0.030]	[0.006]	[0.002]	[0.001]	[<0.001]

Notes. Shown are the results of the long-horizon predictability regressions:

$$r_{t,\,t+H} = \alpha^r(H) + \beta^r(H)\, dp_t + \varepsilon^r_{t,\,t+H},$$

where $r_{t,\,t+H}$ is the H-year continuously compounded log return over time period t until $t+H$, and dp_t is the respective dividend-price ratio at time t. d-p (CRSP) is the dividend-price ratio of all publicly traded corporations, d-p (all-domestic-equity) covers publicly traded and privately held corporations, and d-p (CRSP, adjusted) is the d-p ratio adjusted for composition changes in publicly traded corporations. The estimated slope coefficients $\beta^r(H)$ for the forecasting horizons of one to five years are reported, and their respective p-values for a one-sided test based on Hodrick (1992) standard errors are shown in parentheses; p-values based on a residual resampling bootstrap (BS) are shown in brackets. This table also provides the regression R^2, the bias-corrected R^2_c in percentage terms, and its bootstrapped p-value in brackets. The sample period is 1926–2009 at annual frequency.

return predictability regression for the composition changes-adjusted d-p ratio yields a slope coefficient β^r of 0.86 and an R^2 of 41.9%, compared with a β^r of 0.34 and an R^2 of 13.3% for the conventional CRSP d-p ratio. Moreover, in contrast to the conventional d-p ratio, the slope coefficient and R^2 of the d-p ratio adjusted for composition changes remain statistically significant when bootstrapped p-values are considered.

Recall that the persistence of the all-domestic-equity d-p and adjusted d-p ratio is considerably lower than that of the CRSP d-p ratio (see Table 1). That is, spurious inference regarding the magnitude of the long-horizon R^2 in the OLS estimation is reduced. Even though persistence is weaker, we observe a higher R^2 for these d-p ratios, supporting the notion that long-horizon returns are indeed more predictable.

Overall, the all-domestic-equity and adjusted d-p ratio provide an improvement over the commonly used d-p ratio with respect to several issues. The ratios show no sign of structural breaks and are less persistent. The in-sample prediction results are considerably better and persist when small-sample bias adjustments are taken into account. For long-horizon predictions, the explained return variation is higher, even though persistence of the ratios is lower. Furthermore, the prediction relation is stable over time, and there is evidence of out-of-sample predictability.

6. Further Analyses

6.1. Comparison with Repurchase-Adjusted Ratios

Since the 1970s, traded firms have made increasing use of repurchases (e.g., Grullon and Michaely 2002, Boudoukh et al. 2007). For this reason, Boudoukh et al. (2007) argue that dividends alone might not fully capture cash flows to investors and therefore augment the traditional d-p ratio with repurchases, referring to this as the "total payout yield." However, whether or not the adjustment for repurchases is appropriate is open to debate in the literature. Cochrane (2008) objects that the d-p ratio is not wrong in an accounting sense, as repurchases merely delay the eventual payment of dividends. Koijen and Van Nieuwerburgh (2011) reason that whether a repurchase adjustment is appropriate depends on whether or not one considers an investor who participates in every stock repurchase. In the end, it is an empirical question as to whether the repurchase adjustment is relevant for asset prices (Boudoukh et al. 2007).

For this reason, I compare the all-domestic-equity d-p ratio and the d-p ratio adjusted for composition changes with the repurchase-adjusted d-p ratio regarding their ability to forecast market returns. I update both repurchase-adjusted d-p ratios of Boudoukh et al. (2007)—the measure based on the statement of cash flows [CF] and the one based on changes in treasury stock [TS]—to the end of my sample period. As before, monthly dividends and repurchases are summed up over the year and divided by end-of-year market capitalization.[9] Table 5 shows the correlation matrix of the conventional d-p ratio and the different adjusted ratios. The different ratios are, as expected, highly correlated, which naturally makes it difficult to identify the individual effect of a variable in a multivariate predictive regression.

Table 5 Correlation Matrix of Different Dividend-Price Ratios

	(1)	(2)	(3)	(4)	(5)
(1) d-p (CRSP)	1				
(2) d-p (all-domestic-equity)	0.82	1			
(3) d-p (CRSP, adjusted)	0.70	0.89	1		
(4) d-p (repurchase-adjusted) [CF]	0.67	0.90	0.88	1	
(5) d-p (repurchase-adjusted) [TS]	0.93	0.92	0.81	0.88	1

Notes. Shown are the correlations of different dividend-price ratios. d-p (CRSP) is the dividend-price ratio of all publicly traded corporations; d-p (all-domestic-equity) covers publicly traded and privately held corporations; d-p (CRSP, adjusted) is the d-p ratio adjusted for composition changes in publicly traded corporations; and d-p (repurchase-adjusted) is the d-p ratio adjusted for repurchases, where [CF] refers to the cash flow-based measure and [TS] to the treasury stock-based measure of repurchases. The sample period is 1926–2008 at annual frequency. Correlations for the all-domestic-equity d-p ratio refer to the sample period 1945–2008 because it is only available from 1945 onward.

Nevertheless, Table 6 provides a "horse race" comparison of the different predictive variables using univariate and multivariate forecasting regressions. In addition to p-values based on Newey–West standard errors, the table also provides bootstrapped p-values for coefficients as well as for R^2 values. In the bivariate case, I follow the residual resampling procedure proposed by Pontiff and Schall (1998).[10]

When comparing the explained return variation of the univariate predictability regressions in the postwar sample, we see that the all-domestic-equity d-p ratio and the composition changes-adjusted series outperform the repurchase-adjusted series when predicting returns. In the bivariate forecasting regressions, the explained variation in returns or excess returns increases only slightly when including the repurchase-adjusted ratios, suggesting that the ratios do not possess information independent of each other. In the joint model, the coefficients of the repurchase-adjusted ratios are considerably reduced, sometimes close to zero or

[9] For details on the construction of the repurchase-adjusted d-p ratio, see Boudoukh et al. (2007). I thank Michael Roberts for providing the payout data used in their study on his website (http://finance.warton.upenn.edu/mrrobert/styled-9/styled-13/index.html, accessed February 2, 2012).

[10] Pontiff and Schall (1998) note that for multivariate regressions, bootstrapped p-values are slightly different from conventional p-values in that they are computed under the null hypothesis that all independent variables are unrelated to future returns.

Table 6 Forecast Comparison with Repurchase-Adjusted Dividend-Price Ratios

Panel A: Forecast of Return$_{t+1}$, sample period: 1945–2009								
d-p$_t$ (all-domestic-equity)	0.21 (<0.001) [0.045]				0.29 (0.093) [0.110]	0.21 (0.079) [0.172]		
d-p$_t$ (CRSP, adjusted)		0.21 (<0.001) [0.050]					0.24 (0.104) [0.207]	0.15 (0.042) [0.337]
d-p$_t$ (repurchase-adjusted) [CF]			0.17 (0.009) [0.067]		−0.09 (0.657) [0.739]		−0.04 (0.568) [0.596]	
d-p$_t$ (repurchase-adjusted) [TS]				0.18 (0.002) [0.061]		0.00 (0.511) [0.560]		0.07 (0.236) [0.380]
R^2	14.24	13.60	9.55	11.94	14.76	14.25	13.69	14.25
R^2_c	13.16	12.52	8.47	10.89	11.64	11.14	10.50	11.10
R^2 *p*-value (BS)	[0.005]	[0.007]	[0.027]	[0.012]	[0.018]	[0.021]	[0.026]	[0.022]
Panel B: Forecast of Return$_{t+1}$, sample period: 1926–2009								
d-p$_t$ (CRSP, adjusted)		0.17 (0.003) [0.052]					0.23 (0.076) [0.203]	0.17 (0.052) [0.210]
d-p$_t$ (repurchase-adjusted) [CF]			0.12 (0.059) [0.117]				−0.07 (0.658) [0.648]	
d-p$_t$ (repurchase-adjusted) [TS]				0.11 (0.070) [0.134]				0.00 (0.516) [0.595]
R^2		6.30	3.39	3.48			6.65	6.30
R^2_c		5.67	2.77	2.84			4.78	4.39
R^2 *p*-value (BS)		[0.030]	[0.114]	[0.115]			[0.081]	[0.094]

Notes. This table compares the forecasting performance of the all-domestic-equity dividend-price ratio and composition changes-adjusted d-p ratio with the d-p ratio adjusted for repurchases, where [CF] refers to the cash flow-based measure and [TS] is the treasury stock-based measure of repurchases (for details, see Boudoukh et al. 2007). Reported are the estimated slope coefficients, as well as their respective *p*-values for a one-sided test based on Newey–West standard errors (using one lag) (in parentheses) and based on a residual resampling bootstrap (BS) (in brackets). The regression R^2, the bias-corrected R^2_c in percentage terms, and its bootstrapped *p*-value in brackets are also provided. The sample period is 1926–2009 at annual frequency.

even negative. The coefficients of the all-domestic-equity and composition changes-adjusted ratio, in contrast, are generally reduced by a lesser degree. In some specifications the coefficients of the all-domestic-equity and composition changes-adjusted d-p ratio remain significant at the 10% level when Newey–West standard errors are considered, but not in case of the bootstrap, in which they become insignificant. The high correlation between the right-hand side variables apparently makes it difficult to separate the effects of the individual predictors. Consequently, the all-domestic-equity and composition changes-adjusted d-p ratio are not able to drive out the repurchase-adjusted measures entirely (or vice versa).

In the full sample, return predictability regression the slope coefficients and R^2 values of the repurchase-adjusted ratios are no longer significant at the 10% level (bootstrapped *p*-values), but the composition changes-adjusted ratio remains significant with a bias-adjusted R^2 of 5.67% compared with 2.77% and 2.84% for the repurchase-adjusted ratios. This result is in line with the findings of Koijen and Van Nieuwerburgh (2011), who find no evidence for return predictability despite the repurchase adjustment in the sample period 1926–2009 when cash-reinvested dividends instead of market-reinvested dividends are used.[11]

In summary, the results of the forecasting comparison do not indicate a clear winner, but to some extent, they favor the all-domestic-equity and composition changes-adjusted d-p ratio over the repurchase-adjusted ratios, though admittedly only by a narrow margin.[12] The

[11] It should be noted that the repurchase-adjusted ratios remain significant at the 10% level when excess returns instead of returns are used. The rest of the results are qualitatively the same. In no case do the repurchase-adjusted ratios drive out the all-domestic-equity and composition changes-adjusted d-p ratios or add information about future market excess returns.

[12] I also perform a nonnested forecasting comparison of competing ratios (all-domestic-equity and composition changes-adjusted versus conventional and repurchase-adjusted d-p ratio) in sample as well as out of sample. Following Lettau and Ludvigson (2001), I compare the mean squared error (MSE) of two competing models at a time and report the modified Diebold and Mariano (1995) (MDM) test for forecast encompassing proposed by Harvey et al. (1998). In no case

advantage is most distinct in the full sample when predicting returns, in which there is no evidence for return predictability when adjusting for repurchases but evidence for return predictability when adjusting for composition changes.

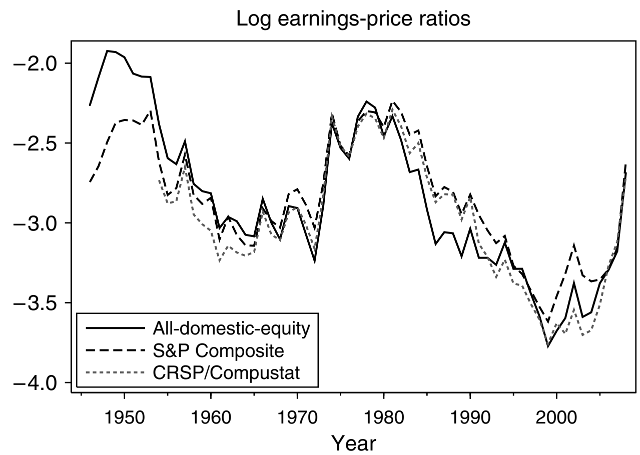
6.2. Comparison of Different Earnings-Price Ratios

Section 4.1 provides evidence that the divergence between the d-p ratio of all corporations and publicly traded corporations is driven by differences in dividends. An alternative explanation for the divergence is a systematic difference in the measurement of the total market capitalization. For example, there might be a mismeasurement in the total market value of all domestic equity. An overly conservative estimate of the market value of closely held firms would reduce the all-domestic-equity d-p ratio and explain the observed pattern. Because the true market value of all (publicly and closely held) equity is unknown, it is not possible to rule out this alternative explanation entirely. However, we can relate the total market value to other firm fundamentals and determine whether there is a divergence, similar to that observed for the d-p ratio. Thus, I compare the log earnings-price (e-p) ratio of all domestic equity against that of the CRSP/Compustat merged sample and, as an additional robustness check, against that of the S&P Composite provided by Schiller (2005). A divergence of the overall e-p ratio and the e-p ratios of publicly traded firms would indicate that the overall market value is systematically under- or overestimated. When comparing these figures, one has to bear in mind that earnings aggregates of NIPA for the all-domestic-equity e-p ratio are not entirely comparable to that of the S&P and CRSP/Compustat merged sample because of different accounting standards.

Nevertheless, we do not observe such a divergence for the e-p ratio in the latter part of the sample, as Figure 4 demonstrates. The relationship of earnings to market value is similar for traded firms and the overall population of all firms, whereas the relationship of dividends to market capitalization differs. There is a slight divergence in the e-p ratios in the early years of the sample, which corresponds to the divergence in d-p ratios in these years, indicating that the market value of all domestic corporations might be underestimated in the early years of the flow of funds accounts statistic. However, from the mid-1950s onward, there is no considerable divergence of the series, suggesting that the difference in dividends rather than in market capitalization is the driver of the divergence in the two d-p ratios.

does a competing ratio have a statistically significant lower in-sample or out-of-sample MSE than the all-domestic-equity and composition changes-adjusted d-p ratio. In the full sample, the composition changes-adjusted ratio has a lower MSE than the conventional and repurchase-adjusted ratio when predicting returns with an MDM test significant at the 10% level both in sample and out of sample.

Figure 4 **Comparison of Earnings-Price Ratios**



Notes. This figure shows the log earnings-price (e-p) ratio of the all-domestic-equity, S&P Composite, and CRSP/Compustat merged data set over time. Earnings are smoothed by taking a five-year moving average. The sample period is 1945–2008 (1955–2008 for CRSP/Compustat) at annual frequency.

7. Concluding Remarks

This paper compares the all-domestic-equity d-p ratio, consisting of publicly traded and privately held firms, against the d-p ratio of publicly traded firms. Although the two d-p ratios follow each other closely from 1945 until about 1970, they substantially diverge afterward. I provide evidence that this divergence can be explained by the systematic composition changes of publicly traded firms. I adjust the d-p ratio for composition changes and use this adjusted d-p ratio and the all-domestic-equity d-p ratio as predictors for stock market returns.

These variables provide improvements on several issues. The all-domestic-equity and composition changes-adjusted d-p ratios are less persistent. In-sample evidence for return predictability is strengthened and is also present after correcting estimates for small-sample bias. The forecasting relation remains stable over time, even during the 1990s, and there is evidence for out-of-sample predictability. Moreover, a forecasting comparison favors the composition changes-adjusted d-p ratio over the repurchase-adjusted ratio when predicting market returns in the full sample. Overall, ignoring the fact that the composition of publicly traded firms changed over time weakens the available evidence for return predictability.

This result points to the importance of a sample selection problem when measuring the time series of financial ratios, which has largely been ignored in the literature thus far. Publicly traded firms represent only a subsample of the entire population of firms subject to change from entering and exiting firms. Therefore, values of a financial variable computed from listed firms only are not necessarily comparable across two points in time. A decline in this variable might be partly due to a genuine decline, but it could also be caused by entering and exiting firms that are systematically different. Considering only the change in the plain variable mistakenly attributes the composition change to the genuine decline, distorting one's inferences. Future research on asset pricing should therefore carefully address the impact of changes in the composition of publicly traded firms.

Supplemental Material

Supplemental material to this paper is available at http://dx.doi.org/10.1287/mnsc.2013.1883.

Acknowledgments

For comments and suggestions, the author thanks the department editor Itay Goldstein, an associate editor, two anonymous referees, Lammertjan Dam, Falko Fecht, Michael Koetter, Martin Lettau, Pratish Patel, Michael Roberts, Maik Schmeling, Marie Scholl, Günter Strobl, Monika Trapp, Michael Weber, seminar and conference participants at the Haas School of Business Finance Ph.D. Seminar at the University of California (UC) at Berkeley, the Konstanz-Tübingen Empirical Finance Seminar, the Centre for Financial Research research seminar at the University of Cologne, the 2012 Annual Meeting of the German Finance Association (DGF), and the 2013 Annual Meeting of the European Economic Association (EEA). This paper was partly written when the author was visiting Haas School of Business, UC Berkeley.

Appendix. Data

All-domestic-equity. The log dividend-price ratio of all domestic equity is calculated by dividing the sum of all dividends paid by domestic corporations (D^{all}) by the total market value (MV^{all}) of all domestic corporations and taking the natural logarithm. Dividend data are from NIPA, and market value data are from FFA. The D^{all} dividends are adjusted for capital gains passed through by mutual funds and mutual funds' interest payments: $D^{\text{all}} = (\text{NIPA Table 7.16, line 30}) + (\text{NIPA Table 7.16, line 31}) + (\text{NIPA Table 7.16, line 35})$.[13] Total market value (MV^{all}) represents all domestic corporate equity issues (total − foreign): $MV^{\text{all}} = (\text{FFA Series FL893064105}) - (\text{FFA Series FL263164103})$.[14]

Publicly listed corporations, CRSP. For the construction of the CRSP dividend-price ratios, I consider all stocks traded on NYSE, AMEX, NASDAQ, and Arca (excluding American depository receipts). $R_t = (D_t + P_t)/P_{t-1}$ is the value-weighted market return including dividends, and $Rx_t = P_t/P_{t-1}$ is the value-weighted return excluding dividends. For each month I compute the level of dividends implied by the return, including and excluding dividends: $D_t = (R_t - Rx_t)MV_{t-1}$. The dividends are then aggregated over the year by summing up monthly dividends. Annual log dividend growth is calculated as $\Delta d_t = \ln(D_t/D_{t-1})$. The log dividend-price ratio is calculated by dividing aggregate annual dividends by end-of-year market value and taking logs: $dp_t = \ln(D_t/MV_t)$. The annual log market return is the log return (including dividends) of the value-weighted market return. The annual log market excess return is calculated by subtracting the log risk-free rate measured by the 90-day treasury-bill rate.

[13] Table 7.16 can be accessed through the Bureau of Economic Analysis website (http://www.bea.gov/iTable/iTable.cfm?ReqID=9&step=1#reqid=9&step=1&isuri=1).

[14] These series can be accessed through the Board of Governors of the the Federal Reserve System website (http://www.federalreserve.gov/apps/fof/DisplayTable.aspx?t=l.213).

References

Ang A, Bekaert G (2007) Stock return predictability: Is it there? *Rev. Financial Stud.* 20(3):651–707.
Bansal R, Yaron A (2007) The asset pricing-macro nexus and return-cash flow predictability. Working paper, Wharton School, University of Pennsylvania, Philadelphia.
Bossaerts P, Hillion P (1999) Implementing statistical criteria to select return forecasting models: What do we learn? *Rev. Financial Stud.* 12(2):405–428.
Boudoukh J, Richardson M, Whitelaw RF (2008) The myth of long-horizon predictability. *Rev. Financial Stud.* 21(4):1577–1605.
Boudoukh J, Michaely R, Richardson M, Roberts M (2007) On the importance of measuring payout yield: Implications for empirical asset pricing. *J. Finance* 62(2):877–915.
Brav A, Graham JR, Harvey CR, Michaely R (2005) Payout policy in the 21st century. *J. Financial Econom.* 77(3):483–527.
Campbell JY (1991) A variance decomposition for stock returns. *Econom. J.* 101(405):157–179.
Campbell JY, Shiller RJ (1988) The dividend-price ratio and expectations of future dividends and discount factors. *Rev. Financial Stud.* 1(3):195–228.
Campbell JY, Thompson SB (2008) Predicting excess stock returns out of sample: Can anything beat the historical average? *Rev. Financial Stud.* 21(4):1509–1531.
Chen L (2009) On the reversal of return and dividend growth predictability: A tale of two periods. *J. Financial Econom.* 92(1):128–151.
Chen L, Da Z, Priestley R (2012) Dividend smoothing and predictability. *Management Sci.* 58(10):1834–1853.
Cochrane JH (2008) The dog that did not bark: A defense of return predictability. *Rev. Financial Stud.* 21(4):1533–1575.
DeAngelo H, DeAngelo L, Skinner DJ (2004) Are dividends disappearing? Dividend concentration and the consolidation of earnings. *J. Financial Econom.* 72(3):425–456.
Dickey DA, Fuller WA (1979) Distribution of the estimators for autoregressive time series with a unit root. *J. Amer. Statist. Assoc.* 74(366):427–431.
Diebold FX, Mariano RS (1995) Comparing predictive accuracy. *J. Bus. Econom. Statist.* 13(3):134–144.
Fama EF, French KR (1988) Dividend yields and expected stock returns. *J. Financial Econom.* 22(1):3–25.
Fama EF, French KR (2001) Disappearing dividends: Changing firm characteristics or lower propensity to pay? *J. Financial Econom.* 60(1):3–43.
Ferson WE, Sarkissian S, Simin TT (2003) Spurious regressions in financial economics? *J. Finance* 58(4):1393–1413.
Goyal A, Welch I (2003) Predicting the equity premium with dividend ratios. *Management Sci.* 49(5):639–654.
Goyal A, Welch I (2008) A comprehensive look at the empirical performance of equity premium prediction. *Rev. Financial Stud.* 21(4):1455–1508.
Grullon G, Michaely R (2002) Dividends, share repurchases, and the substitution hypothesis. *J. Finance* 57(4):1649–1684.
Harvey DS, Leybourne SJ, Newbold P (1998) Tests for forecast encompassing. *J. Bus. Econom. Statist.* 16(2):254–259.
Hodrick RJ (1992) Dividend yields and expected stock returns: Alternative procedures for inference and measurement. *Rev. Financial Stud.* 5(3):357–386.

Inoue A, Kilian L (2005) In-sample or out-of-sample tests of predictability: Which one should we use? *Econometric Rev.* 23(4):371–402.

Kandel S, Stambaugh RF (1989) Modeling expected stock returns for long and short horizons. Working paper, Wharton School, University of Pennsylvania, Philadelphia.

Kendall MG (1954) Note on bias in the estimation of autocorrelation. *Biometrika* 41(3–4):403–404.

Kilian L (1999) Exchange rates and monetary fundamentals: What do we learn from long-horizon regressions? *J. Appl. Econometrics* 14(5):491–510.

Koijen RSJ, Van Nieuwerburgh S (2011) Predictability of returns and cash flows. *Annual Rev. Financial Econom.* 3(1):467–491.

Kothari SP, Shanken J (1997) Book-to-market, dividend yield, and expected market returns: A time-series analysis. *J. Financial Econom.* 44(2):169–203.

Krueger D, Perri F (2006) Does income inequality lead to consumption inequality? Evidence and theory. *Rev. Econom. Stud.* 73(1):163–193.

Larrain B, Yogo M (2008) Does firm value move too much to be justified by subsequent changes in cash flow? *J. Financial Econom.* 87(1):200–226.

Leary MT, Michaely R (2011) Determinants of dividend smoothing: Empirical evidence. *Rev. Financial Stud.* 24(10):3197–3249.

Legel E, Bennett K, Parisi M (2003) The effects of tax reform on the structure of U.S. business. Technical report, U.S. Internal Revenue Service, Washington, DC.

Lettau M, Ludvigson S (2001) Consumption, aggregate wealth, and expected stock returns. *J. Finance* 56(3):815–849.

Lettau M, Ludvigson SC (2010) Measuring and modeling variation in the risk-return trade-off. Ait-Sahalia Y, Hansen LP, eds. *Handbook of Financial Econometrics* (Elsevier Science, Amsterdam), 617–690.

Lettau M, Van Nieuwerburgh S (2008) Reconciling the return predictability evidence. *Rev. Financial Stud.* 21(4):1607–1652.

Lettau M, Ludvigson SC, Wachter JA (2008) The declining equity premium: What role does macroeconomic risk play? *Rev. Financial Stud.* 21(4):1653–1687.

Lustig HN, Van Nieuwerburgh SG (2005) Housing collateral, consumption insurance, and risk premia: An empirical perspective. *J. Finance* 60(3):1167–1219.

Mark NC (1995) Exchange rates and fundamentals: Evidence on long-horizon predictability. *Amer. Econom. Rev.* 85(1):201–218.

McCracken MW (2007) Asymptotics for out of sample tests of Granger causality. *J. Econometrics* 140(2):719–752.

McKinnon JD (2010) When big business enjoys being small. *Wall Street Journal* (eastern ed.) (September 18) A.4.

McGrattan ER, Prescott EC (2005) Taxes, regulations, and the value of U.S. and U.K. corporations. *Rev. Econom. Stud.* 72(3):767–796.

Nelson CR, Kim MJ (1993) Predictable stock returns: The role of small sample bias. *J. Finance* 48(2):641–661.

Newey WK, West KD (1987) A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica* 55(3):703–708.

Paye BS, Timmermann A (2006) Instability of return prediction models. *J. Empirical Finance* 13(3):274–315.

Petrick KA (2002) Corporate profits: Profits before tax, profits tax liability, and dividends. Technical report, Bureau of Economic Analysis, U.S. Department of Commerce, Washington, DC.

Petska T (1996) Taxes and organizational choice: An analysis of trends, 1985–1992. *SOI Bull.* 15:86–102.

Pontiff J, Schall LD (1998) Book-to-market ratios as predictors of market returns. *J. Financial Econom.* 49(2):141–160.

Schiller RJ (2005) *Irrational Exuberance*, 2nd ed. (Princeton University Press, Princeton, NJ).

Stambaugh RF (1999) Predictive regressions. *J. Financial Econom.* 54(3):375–421.

Valkanov R (2003) Long-horizon regressions: Theoretical results and applications. *J. Financial Econom.* 68(2):201–232.

Van Binsbergen JH, Koijen RSJ (2010) Predictive regressions: A present-value approach. *J. Finance* 65(4):1439–1471.

Viceira LM (1997) Testing for structural change in the predictability of asset returns. Working paper, Harvard University, Cambridge, MA.