Buyout Activity and the Equity Risk Premium - Evidence from the U.K.*

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ABSTRACT

We study the relationship between the aggregate risk premium, public-to-private leveraged buyout (LBO) activity, and target betas in the U.K. following recent literature covering the topic in a U.S. setting. In addition, we revisit the U.S. market and relate LBO fundraising activity to variations in the U.S. equity risk premium. Our samples include 1331 U.S. LBOs and 375 U.K. LBOs covering 1982Q4 - 2016Q4 and 1998Q2 - 2016Q4 respectively. Our findings in the U.K. are in line with the evidence documented in the U.S. and further support the notion that variations in the risk premium is a key driver of LBO activity. However, in the U.S. data, we find that the statistical significance of the results is sensitive to how the risk premium is estimated. In particular, a risk premium estimated with an updated version of *cay*, a proxy for the consumption-wealth ratio, leaves most results insignificant. We also find that in contrast to the U.S., target betas in the U.K. are lower when LBO activity is high. Finally, our fundraising data suggests that the decision by limited partners to invest in private equity may also be highly influenced by fluctuations in the risk premium.

Keywords: Private Equity, Buyout Activity, LBO, Equity Risk Premium, *cay*

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WHEN A SPECIALIZED INVESTMENT FIRM finances the acquisition of a public or private company with substantial leverage and limited equity, the transaction is called a leveraged buyout (LBO). Although the term private equity (PE) encompass both leveraged buyouts and venture capital, a private equity firm typically refers to a firm engaged only in the former. This type of acquisition technique first rose in prominence in the U.S. in the 1980s. Private equity as an asset class has grown significantly in size over the last two decades and according to Preqin, a data provider for the alternative assets industry, $453 billion was committed to private equity funds in 2017, the largest amount ever raised in a single year (Preqin, 2018). However, the leveraged buyout market has since its inception displayed a cyclical pattern. Apart from the late 1980s, LBO activity has been particularly vivid in the late 1990s and just prior to the financial crisis of 2008. When trying to explain what drives these booms and busts of private equity, previous literature has mainly focused on credit-centric stories but a recent paper by Haddad et al. (2017) shows that the key driver, like for many other economic phenomena, might in fact be the aggregate risk premium[1].

Haddad et al. (2017) provide two main arguments for the theoretical foundation of the connection between buyout activity and the aggregate risk premium. First, when the risk premium is high, future gains of a buyout deal are discounted at a higher rate thereby making the investment less attractive. Second, concentrated illiquid positions, such as privately held companies, are particularly unattractive to investors when the risk premium is high. The authors argue that these two forces predict that buyout activity is negatively correlated with the risk premium.

Based on the assumption that both limited partners and general partners must benefit in the buyout transaction, Haddad et al. (2017) construct a model that relates the buyout decision to the aggregate risk premium. The authors outline two key mechanisms. First, a so-called performance channel which incorporates potential performance improvements valued using a Net Present Value rule. With a higher risk premium, future performance gains are discounted at a higher rate which lowers valuation and leads to fewer buyouts being made. Second, an illiquidity channel, resulting from the separation between the agent (GP) and the principal (LPs), dictates the cost of compensating the general partner for bearing excessive risk. The general partner has to bear excess risk to be motivated to implement the changes in the acquired firm. Therefore, when the risk premium decreases the excess risk also decreases and compensating the general partner becomes less costly[2]. The authors derive the following performance and illiquidity inequality:

$$(p_H - 1)(\mu - \beta \bar{R}^e_m) \geq \frac{1}{2}\bar{R}^e_m \frac{1}{\theta^*_m}(k^*_1 p_H \beta - \theta^*_m)^2 + \frac{1}{2}\gamma\sigma_i^2 k^*_1 {p_H}^2, \tag{1}$$

where β is the equity market beta, p_H is a factor of how much the firm output is increased if a buyout deal occur through operational improvements, R^e_m is the expected market return, μ is the average output of the firm, θ^*_m is the general partners position in the public market, k^*_1 is a component controlling the riskiness of the targets variable output, γ is the risk aversion of the

[1] For example, for the risk premium's effect on variations in firm investments see Berk et al. (1999).

[2] In essence, the magnitude of the risk premium reflects the rate of return required for bearing a certain amount of risk.

general partner and σ_i^2 is the idiosyncratic volatility of the buyout target.

The finalized model states that if there is a positive return net of the general partners compensation, a buyout deal occurs. In other words, if the performance channel (the left-hand side of Equation 1) exceeds the illiquidity channel (the right-hand side of Equation 1), a buyout deal occurs.

The model leads to a set of testable predictions[3]. First, buyout activity should be lower in times of a high risk premium. Second, a firm is more likely to be acquired if the firm either has a low market beta or a low idiosyncratic risk and the beta consideration is especially important when the risk premium is high. In other words, target betas should be lower on average when the risk premium is high. Haddad et al. (2017) find strong empirical support for both predictions when studying a sample of U.S. public-to-private LBOs.

Haddad et al. (2017) study deal activity and the equity risk premium up until 2011. Since then, the equity risk premium should in expectation decreased following an exceptionally low interest rate environment (Campbell, 1987). Thus, we would expect the number of private equity deals to have increased. However, the number of deals in recent years has been at a modest level with aggregated values invested in 2017 not even reaching half of the value invested in 2007, the peak year of private equity activity (Preqin, 2018). As noted, the unexpectedly low number of deals has not been a result of less capital committed to private equity firms by limited partners. The low number of deals coupled with the high levels of fundraising has brought the dry powder level to a record high. The potential change in the relationship between the equity risk premium and deal activity calls for further analysis.

In this paper we begin by revisiting the results on public-to-private LBOs[4] found by Haddad et al. (2017) in the U.S. market to test if the results still holds true in the light of unprecedented levels of dry powder. We then turn to the U.K. LBO market to test if the aggregate risk premium explains as much of LBO activity as it seems to do in the U.S. Apart from the U.K. being the largest LBO market in Europe, it provides an interesting testing ground since LBO activity itself is not as clear-cut in this market as it is in the U.S. The U.K. saw an unrivalled peak in LBO activity in the late 1990s in terms of number of transactions. However, when looking at the value of deals, the peak can instead be found in the in the 2006 - 2007 period (Renneboog and Vansteenkiste, 2017). Following Haddad et al. (2017), we also test if target betas display the same sensitivity to the risk premium and LBO activity as they do in the U.S. We estimate the risk premium by regressing future market returns on a set of factors found to possess predictive capabilities in the existing body of literature. Since forecasting returns itself is a widely covered topic of research, we find it relevant to incorporate some variations in the estimation methods used. We make a distinction between risk premiums estimated out-of-sample (OOS) without look-ahead bias and risk premiums estimated in-sample (IS).

The buyout decision model suggested by Haddad et al. (2017) is based on the assumption that

[3] The two predictions mentioned here are not the only ones tested empirically. For a full list please refer to Haddad et al. (2017)

[4] Public-to-private LBOs are in this paper hereafter referred to simply as LBOs

general partners and limited partners assess each investment in conjunction. In other words, the willingness to invest in any given deal does not only depend on the general partner's assessment of the investment opportunity, but also the limited partner's. One could argue that although this might be true in many cases, a common trait of the private equity industry is the construction of private equity funds where fundraising, i.e. the decision by limited partners to invest, precedes the assessment by general partners of potential investments opportunities. The risk premium might very well still impact the decision to invest by limited partners. One might for instance hypothesize that an institutional investor is more willing seek alternative investment classes, such as private equity, when the interest rate is low or when the stock market is expected to deliver below-average future returns. However, if the risk premium is time-varying, the magnitude of it at the time limited partners choose to invest in the private equity fund might differ from when the general partner decides to acquire a given target. We find this separation interesting and we therefore relate the aggregate risk premium to variations in private equity fundraising for the U.S. market.

We find that the relationship between LBO activity and the aggregate risk premium still holds true in the U.S. although adding the post-2011 data weakens the statistical significance of the results. Furthermore, if estimating the risk premium with a slightly different version of *cay*, one of the input variables used when estimating the risk premium, the results become statistically insignificant. For the U.K., we find support in the data for a strong link between the risk premium and LBO activity, indicating that the findings made by Haddad et al. (2017) are not exclusive to the U.S. market. When regressing U.K. LBO activity on U.K. risk premiums the results are significant at the 1% level. When it comes to the relationship between LBO activity and target betas, we find an interesting contrast in the U.K. data. Here, the relationship between LBO activity and target betas appears to be reversed and statistically significant at the 1% level. Higher number of LBOs, scaled by the number of public firms each quarter, is in contrary to expectations found to be associated with lower target betas. Log transaction value, an alternative way of defining LBO activity, yields results more similar to expectations but the risk premium behaves consistent with the scaled number of LBOs measure; a higher risk premium is found to be associated with higher target betas. In both markets we find that betas have increased over time giving rise to the question if LBO firms behave differently in more recent years when it comes to deal sourcing. Finally, we find that private equity fundraising in the U.S. might also be explained by the aggregate risk premium adding insights to the existing body of literature why institutional investors turn to private equity. The available data limits us from performing the same tests on the U.K. market. This, among other questions raised in this paper leaves a few interesting avenues for future research.

This paper is organized as follows. Section I provides an overview of the existing literature on the topic, Section II details our data sources, sample selection criteria and the methodology of the statistical tests. Section III presents descriptive statistics for our U.S. and U.K. samples. In Section IV the regression results are presented, while we interpret our results and consider their implications in Section V. Finally we conclude the paper and comment on topics for future research in Section VI.

I. Literature Review

A. Development of the Private Equity Industry

Kaplan and Strömberg (2009) show that U.S. private equity fundraising and transaction values as a percentage of the U.S. stock market value has varied over time peaking in the late 1980s, the late 1990s and in 2007. The number of global leveraged buyout transactions has also increased dramatically over the same period. Although the number of private equity transactions has been far below the peak levels of buyout activity in 2006 and 2007, it seems like the private equity industry has recovered quite well since the financial crisis of 2008 in terms of fundraising (Preqin, 2018).

Kaplan and Strömberg (2009) track buyout activity and private equity fundraising as a percentage of the total U.S. stock market value over time and find that fundraising mirrors transactions. The authors note that this pattern is consistent with the hypothesis that activity is determined by systematic mispricings in the debt and equity markets. When the cost of debt is relatively low compared to the cost of equity, private equity firms can take advantage of the difference in an arbitrage fashion.

Renneboog and Simons (2005) describe factors, other than systematic mispricings, that are likely to have contributed to the booms and busts of private equity. Bankruptcies of firms acquired in the U.S. boom in activity during the 1980s caused public and political resistance followed by anti-takeover legislation. Adding a credit crunch and a crisis in the high yield bond market caused the activity to decrease substantially. The surprisingly low activity during the first half of the 1990s could, according to Kaplan (1997), arguably be a result of corporations at the time institutionalizing the focus on shareholder value brought by private equity firms. Renneboog and Simons (2005) argue that the increase in activity in the U.S. during the late 1990s was a result of increased costs associated with being listed at a stock exchange due to the Sarbanes-Oxley act. Smaller firms were affected particularly strong since the cost of adhering to the regulation were largely fixed.

With the exception of the U.K., Wright et al. (2006) conclude that European LBO activity did not materialize until the mid-1990s. They state that one of the reasons why large sized LBOs increased in number in the late 1990s was the introduction of a European subordinated high yield debt market, able to fund buyouts. Previously, funding for large European LBOs had to rely on the U.S. high yield bond market. The authors also outline a number of phases in the U.K. buyout market. Just like in the U.S., U.K. buyout activity took off in the early 1980s, although in a much smaller scale. The type of targets in the early phase were mostly distressed companies affected by the 1979-1982 U.K. recession. From the mid-1980s to the end of the decade, activity increased and corporate refocusing became a more and more common investment rationale. Following the recession of the early 1990s, distressed firms again became more predominant among transaction targets and many of the foreign banks that had helped activity increase in the late 1980s, abandoned the buyout market. The number of transactions recovered during the 1990s but once again saw a sharp drop following the burst of the dot.com bubble.

Renneboog and Vansteenkiste (2017) note that the wave of U.K. public-to-private LBOs, fueled by cheap funding available through the emergence of the securitized debt market following the dot.com bubble, came to a halt when the CDO markets collapsed in late 2007. The authors also outline the variations in U.K., U.S. and continental public-to-private LBOs in terms of volume and value. The number of deals and value of deals each year correspond quite well for both the U.S. and continental Europe. In the U.K. however, there is a clear peak in number of transactions in the late 1990s/early 2000s while the value of transactions is more modest, instead peaking in 2006-2007.

Apart from variations in buyout activity over time, Strömberg (2008) looks at a large dataset, from 1970 to 2007, of global LBOs. The author notes that LBOs have occurred in a wide range of industries, even early in his sample period. He further notes that some trends can be discerned. For instance, the fraction of LBOs carried out in the retail sector dropped from 14% in the 1970s and 1980s to 6% in the 2000s while high-growth sectors such as biotech and computers have increased substantially in more recent decades. The author offers two potential explanations. The changing industry mix could simply be a result of a change in the industry mix of the economy as a whole. On the other hand, it could reflect a deliberate shift, away from the traditional private equity targets operating in mature industries characterized by stable cash flows and high debt capacities.

B. Private Equity and Systematic Risk Exposure

The literature studying private equity-owned firms' systematic risk exposure have mainly focused on estimating betas during private equity firms' holding period. For instance, Axelson et al. (2014) examine a sample of buyout deals from a large fund-of-funds investing in private equity and find a levered beta of 2.2 - 2.4. The authors note that most other studies such as Franzoni et al. (2012) have found considerably lower betas, around 1.0, which is puzzling since it would imply that private equity firms are able to acquire firms with average equity betas, subsequently increasing their leverage dramatically while keeping their systematic risk exposure intact. Ljungqvist and Richardson (2003) find that publicly traded firms, comparable to a set of private portfolio companies, have betas around 1.0 when studying a sample of buyout transactions provided by a large U.S. institutional investor.

However, there are some evidence indicating that private equity firms on average acquire companies with equity betas below 1.0. For example, Frazzini and Pedersen (2014), find an average ex-ante beta below 1.0 with statistical significance when looking at public-to-private LBO deals between 1963 - 2012. The authors suggest that this finding is in line with their hypothesis that agents with more relaxed funding constraints will invest in low beta assets and apply leverage in order to achieve a greater return. They hypothesize that high beta stocks underperform since investors with leverage constraints, such as mutual funds and retail investors, overweight high beta assets in their portfolios, causing those assets to yield lower returns. However, no explicit claim is made that private equity firms in general use a so-called betting-against-beta strategy when investing. It is possible that low beta firms are simply perceived as being less sensitive to the business cycle and thereby more capable of handling the high levels of debt private equity firms employ.

Although the evidence is somewhat mixed, several studies, such as Davis et al. (2014) have found empirical support for the notion that operational improvements in portfolio companies contribute to private equity returns. Given the option to use financial engineering in combination with operational and governance engineering, perhaps beta considerations are less important in deal sourcing from a pure return perspective. However, Guo et al. (2011) find that operating improvements in buyouts carried out between 1990 and 2006 seem to be smaller than their 1980s counterparts. Coupled with the well documented and persistent low beta anomaly first noted by Black (1972) and Black et al. (1972), and more recently by e.g. Frazzini and Pedersen (2014), one might hypothesize that targets' market beta play a role in explaining private equity excess returns.

Haddad et al. (2017) argue that low beta firms are more attractive private equity targets, and also find that the average beta in their public-to-private LBO sample is below the average of non-LBO targets. Furthermore, the authors find that betas on average are higher when the aggregate risk premium is low and LBO activity is high. They argue that since riskier firms have higher cost of capital, greater illiquidity costs and are more sensitive to changes in the risk premium, high beta firms will be particularly weak buyout candidates when the risk premium increases.

C. Deal Sourcing and Time-Varying LBO Performance

Haddad et al. (2017) emphasize that their model and predictions rely on the assumption that the performance improvements private equity firms aim to realize in buyout deals are valued using a Net Present Value rule. When the risk premium increases, and thus the cost of capital, targets are valued lower and fewer deals are made. This assumption stands in contrast to what Gompers et al. (2016) find regarding private equity firms' valuation methods. In their survey, fewer than 20% of the respondents use Adjusted Present Value or WACC-based DCF models to evaluate target firms. The average respondent ranked the NPV method 2.8 out of 10.0 when asked which valuation methods they rely most on. The Internal Rate of Return metric was ranked 9.2 out of 10.0 on average. The authors conclude that it appears like private equity investors do not use NPV or DCF valuation techniques frequently. Nevertheless, the authors note that private equity firms seem to incorporate factors related to systematic and non-systematic risk when determining hurdle rates. Thus, jumping to the conclusion that the buyout decision model constructed by Haddad et al. (2017) is flawed would be unwarranted. It is possible that NPV techniques have been used more extensively in the past (the survey conducted by Gompers et al. (2016) took place in 2012) and private equity firms might implicitly incorporate the current risk premium in one way or another when evaluating deals.

Axelson et al. (2013) find a negative relationship between fund returns and leverage. This supports the theory of Axelson et al. (2009) that the higher leverage used by private equity firms in hot markets is not necessarily in the best interest of the limited partners investing in their funds. Axelson et al. (2009) argue that during recessions, even the valuable investment opportunities that exist will be difficult to finance. In contrast, during market booms the easy access to leverage enables private equity firms to finance even invaluable projects. In other words, when leverage is easily accessible, private equity firms seem to overpay for deals. Empirical evidence implying a similar conclusion is found by Kaplan and Schoar (2005). They observe that funds raised in boom times are less likely to raise follow-on funds.

D. Forecasting Returns and the Equity Risk Premium

A large body of literature covers factors that have forecasting capabilities of future equity returns. For instance, Fama and French (1988) find that the dividend-price ratio is a relatively strong predictor of future stock market returns. They note that the predictable component of stock price variation is small for shorter time horizons, but often explain more than 25% of variations in two to four year horizons. Campbell (1987) finds that the term structure of interest rates predicts stock returns. His findings suggest that uncertainty about short-term nominal interest rates is an important factor in the pricing of long-term assets.

Lettau and Ludvigson (2001) find that the consumption-wealth ratio, or *cay*, has predictive power and outperform the dividend-price ratio at shorter horizons. Brennan and Xia (2005) critique the *cay* measure and suggest that its predictive power stems from look-ahead bias in its construction. They find that if the *cay* measure is estimated out-of-sample (OOS), it loses its predictive power. Cochrane (2011) argues that *cay* performs well when forecasting one-period returns, capturing wiggles in business cycle frequency, while not affecting the overall trend and performs much worse in long-run forecasts. As emphasized by Welch and Goyal (2008), plenty of models shown to predict market returns perform much worse or fail entirely out-of-sample. Furthermore, they note that most factors that have been shown do possess predictive capabilities especially fail in more recent decades. Predicting equity returns with only the information available at the time is of course a difficult, if not impossible, task. Both the dividend-price ratio and *cay* receive critique from Welch and Goyal (2008) who question their usability in practice.

II. Data and Methodology

A. Defining Leveraged Buyouts

Data on U.K. and U.S. LBOs is collected from the Thomson Reuters SDC Platinum database. Following Haddad et al. (2017), we select completed, 100% acquired, public-to-private deals classified as a Leveraged Buyout or Management Buyout. As pointed out by Officer et al. (2010) among others, the LBO and MBO flags are not completely comprehensive and miss some private equity transactions. Haddad et al. (2017) supplement their sample with target firms acquired by private, financial acquirers where the deal is made for investment purposes. Although the procedure may sound straightforward, it is not entirely clear which deals are added. A more comprehensive sample selection procedure is described in an earlier version of their paper (Haddad et al., 2013). We follow that approach and include completed deals by private acquirers, where the target is 100% acquired and the acquirer is described as an investor group, a financial acquirer or a management group. We exclude any of the above deals described as a spinoff, divestiture or bankruptcy or where the acquirer and target share the same Fama-French 48 industry[5]. The announcement dates recorded in SDC determine the timing of the transaction. We classify any deal where the last available data on debt is six quarters prior to the announcement date as having missing accounting data.

For the U.S. we follow Haddad et al. (2017) and start our study in 1982Q4, the point in time where LBOs started to become more frequent. The emergence of LBO activity in the U.K. lagged the U.S., becoming much more frequent in the late 1990s. As a result, we start our U.K. sample in 1998Q2. Both samples end in 2016Q4. For the U.S., the sample comprises 1331 LBOs. For the U.K., the sample is comprised of 375 LBOs. When estimating the targets' ex-ante betas, the samples shrink to 1007 for the U.S. and 279 for the U.K. mostly due to missing or infrequent price data or missing accounting data in CRSP, COMPUSTAT or Datastream.

B. Activity

In line with Haddad et al. (2017), we scale the number of LBOs each quarter by the number of listed firms in the U.S. and the U.K. respectively. The number of U.S. firms are gathered from CRSP and the data is available at a quarterly basis. For the U.K., we collect the data from the historical records of the number of listed firms available on the London Stock Exchange website. The data is only available on a yearly basis. We therefore assign a value to the quarters missing data by interpolation for the first, second and third quarter each year[6]. Although the data for the U.K. sample lacks the precision of its U.S. counterpart, we believe it is sufficient to capture the variation in number of firms over time. The quarterly changes in the U.S. sample are mostly small and correspond closely to a time series of yearly observations with quarterly interpolations.

As mentioned in Section I.A, Renneboog and Vansteenkiste (2017) note that the peak in number of deals and the peak in target transaction values did not occur at the same time in the U.K. To the contrast, the two measures generally follow each other more closely in the U.S. and yields equivalent regression results in the tests by Haddad et al. (2017). Therefore, we also construct an activity measure based on transaction values for our U.K. sample but not for the U.S. sample. The data on transaction values is collected from the SDC platinum database. The database does not provide transaction values for 15 out of the 375 deals. We define the measure in logarithmic terms to reduce skewness.

C. Estimating the Risk Premium

The risk premiums are estimated using OLS regressions where the dependent variable is the next three years' annualized return on the value-weighted market portfolio in excess of the three-month T-bill rate on a quarterly basis. For the U.S. the annualized return is calculated on the CRSP

[5] The classifications can be found on Kenneth R. French website, `http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/Data_Library/det_48_ind_port.html`

[6] The first quarter each year is two thirds of the number of firms in the end of the previous year and one third of the number of firms in the end of the year. The second quarter is the average between the end of the year and the end of the previous year. The third quarter has a two thirds weight on the number of firms at the end of the year and one thirds weight on the end of the previous year.

NYSE/AMEX/NASDAQ/ARCA index. For the U.K. the annualized return is calculated on the Datastream Total Market U.K. Equity index. The explanatory variables have been shown to predict market returns in previous studies, as outlined in Section I.D, and include the dividend-price ratio (D/P), the three-month T-bill yield and *cay*, introduced by Lettau and Ludvigson (2001). The dividend-price ratio is obtained using CRSP NYSE/AMEX/NASDAQ/ARCA data on monthly returns for the U.S. sample. For the U.K., we use the dividend-price ratio from the Datastream Total Market U.K. Equity index. The risk-free rate used for the U.K. is the Thomson Reuters U.K. Three-month T-Bills Bid Yield, available in Datastream. The risk-free rate used in the U.S. sample is the three-month T-bill annualized yield-to-maturity available from CRSP.

Since the study conducted by Haddad et al. (2017), Lettau and Ludvigson have revised the *cay* measure. They now construct *cay* using personal consumption expenditures (PCE) instead of nondurables and services (NDS). As emphasized in the *cay* revision notes (Lettau and Ludvigson, 2015), total consumption is unobservable and data on expenditures used to proxy for consumption is not necessarily accurate. The authors point out that if expenditures on nondurables in relation to total consumption is constant over time, omitting expenditures on durables will not affect the results. The previous choice of excluding durables depended on this assumption. However, the fraction of expenditures on NDS to PCE has decreased quite substantially over the past decades making it difficult to ignore the possibility that PCE better captures total consumption.

For the U.S., the PCE measure of *cay* is available through Martin Lettaus website[7] where log consumption, labor income and asset wealth data is also provided. We will return to the effect of using the PCE measure of *cay* instead of the NDS measure in Section IV but given the somewhat arbitrary choice between PCE *cay* and NDS *cay*, we also construct the NDS *cay* measure using data from the U.S. Bureau of Economic Analysis via the Federal Reserve Bank of St. Louis. For the U.K. sample we also construct a PCE and a NDS version of *cay* using data from the Office for National Statistics. We leave a more thorough description of the data sources used to construct the U.K and U.S. *cay* measures to Appendix C and Appendix D. Following Stock and Watson (1993) and Lettau and Ludvigson (2001) we estimate *cay* using a Dynamic Least Squares regression, where the following equation is estimated:

$$c_t = \alpha + \beta_a^s * a_t + \beta_y^s * y_t + \sum_{i=-k}^{k} b_{a,i}^s * \Delta a_{t-i} + \sum_{i=-k}^{k} b_{y,i}^s * \Delta y_{t-i}, \quad t = k+1, \ldots, +s-k, \tag{2}$$

where c is aggregate consumption, a is aggregate wealth and y is aggregate income. k is the number of leads/lags. Following Lettau and Ludvigson (2001) we choose 8 leads/lags for the U.S. sample. Publicly accessible quarterly data on the input variables for the U.K. is not available until 1987. Due to the significantly shorter estimation window we instead choose k=1 in line with Della Corte et al. (2010), who also constructs *cay* in a European setting. We comment on the potential biases the shorter estimation window may give rise to in Section V. The estimated coefficients

[7] `https://sites.google.com/view/martinlettau/data`

provide us with the *cay* measure as follows:

$$cay \equiv \hat{cay}_t = c_t - \hat{\beta}_a * a_t - \hat{\beta}_y * y_t, \quad t = 1, \ldots, T-k. \tag{3}$$

To make sure that our methods of constructing NDS, and thus NDS *cay*, are robust we estimate PCE on our own and reconstruct an equivalent PCE *cay* measure to the one available through Martin Lettaus website. Our measure of PCE has a correlation close to 1.0. Using our measure of PCE together with Lettaus values for asset wealth and labor income consequently results in a measure of *cay* that has a correlation close to 1.0 with the *cay* supplied by Lettau[8].

As mentioned in Section I.C, an underlying assumption necessary to rationalize the relationship between the aggregate risk premium and LBO activity as in Haddad et al. (2017), is that private equity firms use the risk premium to evaluate deals, which might very well be the case. How these private equity firms estimate the risk premium is of course unknown but they would clearly only be able to do so with the information available at the time. Therefore, one could argue that between the choice of either a risk premium estimated on a rolling basis, with information that would have been available at the time, or a risk premium estimated over the whole sample period, the former would be more appropriate. However, by using all currently available data, the estimation of the true relationship between future returns and our set of explanatory variables becomes more accurate. Since we do not know how the average private equity firm would estimate the risk premium we include risk premiums estimated out-of-sample and risk premiums estimated in-sample in our analysis.

The out-of-sample risk premium estimations give rise to some further considerations. As pointed out in Section I.D, Brennan and Xia (2005) note that the *cay* measure is itself estimated with look-ahead bias. We take this into consideration and re-estimate the *cay* measure each quarter with the log consumption, income and wealth data. In the risk premium regressions, we assume that a practitioner would update her beliefs not only regarding the current *cay* but also the past values of *cay* as longer and longer horizons are used to estimate it, thus increasing its precision. Consequently, past *cay* values are updated in each risk premium regression.

Since *cay* requires a certain number of data points before a first estimate can be found, we run our out-of-sample regressions on a quarterly basis from 1965Q3 to 2013Q4 for the U.S. sample and 1990Q1 to 2013Q4 for the U.K. sample. In the out-of-sample estimations, a given quarter's risk premium is calculated as follows:

$$E\left(R^e_{M,t+1}\right) = \alpha + \beta_{dp,t-1}(D/P)_t + \beta_{cay,t-1}cay_t + \beta_{tbill,t-1}(\textit{T-Bill})_t, \tag{4}$$

where the D/P ratio, *cay* and T-bill is the then currently available data and the regression coefficients are the most recently available. Since we forecast three years of future returns, at any given point in time, the most recently available coefficient estimates are based on a regression using

[8] The actual correlation between the two PCE consumption measures is 0.999998 while the actual correlation between the two PCE *cay* measures is 0.9998389.

the inputs available three years earlier. For example, in 1998Q1 the left hand side of Equation 4 is the annualized future three years' return for 1995Q1, forcing a practitioner to use *D/P*, *cay* and T-bill data from 1995Q1 and prior. However, the regression coefficients used to construct *cay* has been updated every quarter in-between 1995Q1 and 1998Q1. In the 1998Q1 regression, all past values of *cay* are therefore estimated with the most current available data. The value of *cay* assumed to be known in 1995Q1 is thus updated in 1998Q1, as opposed to the T-bill and *D/P* ratio.

For the in-sample approach we run the regression from 1952Q1 to 2013Q4 for the U.S. (and also between 1954Q1 to 2010Q3 for comparability with the risk premium used by Haddad et al. (2017)) and 1987Q1 to 2013Q4 for the U.K. For full comparability, the U.S. risk premium estimated between 1954Q1 to 2010Q3 is based on the *cay* estimates that were provided in 2011Q4 on Martin Lettau's web page which was then based on the NDS measure. Re-estimations of asset wealth and labor income by the Bureau of Economic Analysis results in an NDS *cay* that differ slightly if estimated up until 2011 today. Regarding the timing of the data, we use the risk premium for the previous quarter (estimated the last month of the quarter) when running regressions on LBO activity and target betas. The same applies to the credit factors described below.

D. Credit Factors

For the U.K. sample, we test LBO activity on both the risk premium and two credit market factors, namely the EBITDA spread (Strömberg, 2008). The HY spread[9]. The EBITDA spread is defined as the median EBTIDA/EV less the Merrill Lynch High Yield and the HY spread is simply the Merrill Lynch high yield less the risk-free rate. For the U.S. sample we also include the GZ spread (Gilchrist and Zakrajšek, 2012) in addition to the HY spread and the EBITDA spread. The GZ spread is an average credit spread on senior unsecured bonds issued by non-financial firms and has shown to predict economic activity. Data on the GZ spread is available through Simon Gilchrist's website[10] but a U.K. equivalent is unfortunately not available. However, as we will show in Section IV, the GZ spread is not found to explain much of variations in LBO activity in the U.S.

For the U.S. sample, EBITDA and enterprise value data for all public firms are collected from COMPUSTAT. For the U.K. sample, EBITDA and Enterprise value data are collected from Datastream where we consider all constituents in the FTSE All Share Index each quarter. The risk-free rate used in the U.S. sample is the three-month T-bill rate available at the Federal Reserve Bank of St. Louis' website. The same website is used to collect data on the Merrill Lynch High Yield Master II Bond Index for the U.S. and the Merrill Lynch Euro High Yield Bond Index for the U.K. Data on the Merrill Lynch High Yield Master II is not available before 1997 and the Merrill Lynch Euro High Yield is not available before 1998. We are therefore forced to test the effect on LBO activity from the credit measures based on these yields on a subset of our U.S. sample.

[9] The HY spread is introduced in Haddad et al. (2017) and inspiration comes from Axelson et al. (2013) who find that a similar spread, the Merrill Lynch High Yield Index less LIBOR is correlated with LBO activity

[10] `http://people.bu.edu/sgilchri/Data/data.htm`

E. Betas

It is far from obvious what the optimal return observation frequency is when estimating a firm's market beta. Hawawini (1983) notes that in general, securities with smaller than average market value will have betas decreasing in magnitude as the return interval shortens and vice versa for larger than average market value firms. The difference can in some cases be dramatic between daily and monthly return frequencies[11].

For comparability we follow Haddad et al. (2017) when estimating betas. They use monthly return data and although no explicit motivation is given regarding this choice, a potential explanation is reducing the noise daily or weekly return data can entail. Another consideration is observation window where we use two years of monthly returns. A longer window might increase the precision of the beta estimate but might at the same time capture a time period that is not representative of the firm's market risk exposure at the deal announcement. Haddad et al. (2017) find that LBO targets are more levered than equivalent public firms on average. Since the equity beta is dependent on a firm's financial leverage, we unlever the betas in order to estimate total firm risk. We assume a debt beta of zero and a tax rate, τ, of 35% for the U.S. and a tax rate, τ, of 30% in the U.K. reflecting the differences in corporate tax rates between the two countries during our regression windows. Thus, the unlevered beta is

$$\beta_U = \beta_L \frac{1}{1 + (1-\tau) * \frac{Debt}{mktcap}}. \tag{5}$$

Stock price and accounting data used to calculate unlevered betas are collected from CRSP and COMPUSTAT for the U.S. sample and Thomson Reuters Datastream for the U.K. sample. Since we unlever the target betas with the assumption of a debt beta equal to zero, we trim the top 5% D/E-ratio targets for the U.S. sample. For these companies, a debt beta of zero is an unrealistic assumption. For the U.K. we find that the D/E-ratios are lower overall and we therefore remove the companies that surpassed the threshold for the top 5% D/E-ratio targets in the U.S. We also follow Haddad et al. (2017) and trim the top and bottom 5% beta targets in our U.S. sample to reduce the impact of large outliers. For the U.K., we find that it is sufficient to trim the top and bottom 2.5% to get a similar dispersion in betas as in the U.S. (see Table I and II).

F. Private Equity Fundraising

We collect data for all buyout funds available in Thomson Reuters Eikon private equity and venture capital database. We select funds defined as either Buyout, Other Private Equity or Generalist Private Equity categorized as being located in the Americas. According to Metrick and Yasuda (2010) the average number of investments in a buyout firm is 14.76 while the median is 12. The Thomson Reuters data in its original form includes several funds with substantially more deals attached to them. Funds that have more than 20 deals are therefore excluded in our sample.

[11] For one particular firm, Hawawini finds a 53% difference in beta value when comparing daily and monthly return frequencies.

Although the cut-off point is somewhat arbitrary, a majority of the funds that are excluded are large and classified as an unspecified fund of a given private equity firm. It is likely that these unspecified funds are in fact aggregated from a variety of smaller unknown funds. Including these unspecified funds would thus allocate more fundraising to a specific quarter than what is actually the case. We also exclude funds where the difference between the last investment and the first investment exceeds 12 years. The data does not specify in which quarter of a year the fund was closed. We therefore let the first investment date of the fund act as a proxy for the date when the fund was raised. Since the fundraising will have taken place before the first investment we run our regressions with a variety of lags on the risk premium. The size of the funds in a given quarter is scaled by the corresponding market value in the CRSP combined NYSE/AMEX/NASDAQ index.

The Thomson Reuters data lacks fund location on a country level. Although it is reasonable to assume that the funds categorized as being located in the Americas mostly refers to U.S. funds, assuming that most funds located in Europe refers to U.K.-based funds is more far-fetched. Given the additional impreciseness of the data in terms of fund closing dates, we choose to exclude the U.K. LBO market in our fundraising analysis.

G. Treatment of Seasonality, Heteroscedasticity and Serial Correlation

Given the persistence in our independent variables in the regressions, we apply Newey-West standard errors lagged over the prior four quarters. Newey-West standard errors also account for potential heteroscedasticity. This is in line with previous literature and used by e.g. Haddad et al. (2017). Newey-West standard errors are applied in all regressions in Section IV while quarterly dummy variables are included in all regressions, to account for seasonality in the buyout activity measure, except when a measure of beta is the dependent variable.

III. Descriptive Statistics

A. U.S. Sample Summary Statistics

In Table I, we present summary statistics for our U.S. sample. There are 9.72 LBOs taking place per quarter on average. The scaled activity measure shows that 0,138% of public firms on average are taken private per quarter. Transaction values, asset values and enterprise values are collected from the SDC Platinum database. The corresponding quarterly values are \$9.74 bn, \$10.15 bn and \$11.5 bn.[12] All three measures have large standard deviations compared to their means[13].

Summary statistics for the aggregate factors, both credit factors and the risk premiums, can also be found in Table I. It is interesting to note that there is some differences in regards to standard deviation, minimum and maximum values between the two in-sample NDS *cay* risk premiums

[12] 92 of the 1331 targets in the sample lack transaction value data. 119 of the 1331 targets in the sample lacks asset value data. 200 of the 1331 targets in the sample lacks enterprise value data.

[13] This, and the fact that the values at the 75th percentile suggests that there is a positive skew, indicates that these values should be logged if they were to be used in a regression.

resulting from the different estimation periods. In Figure 1 one can see that the maximum and minimum values does not occur after 2011Q4. The difference in minimum and maximum values are therefore entirely up to the difference in regression coefficient outputs. Regression coefficients for our in-sample risk premium estimations are reported in Appendix B.

Unlevered betas in our U.S. sample are on average 0.79 while the average market beta is 1.01 for the 827 deals included in our regressions. The unlevered betas are, as should be expected, on average very similar to the ones in the sample of Haddad et al. (2017).

Table I
Summary Statistics, U.S.

This table presents quarterly summary statistics for U.S. LBO activity, credit factors, accounting measures, transaction values and risk premiums. It also includes the Equity β and its corresponding Unlevered β. Number of LBOs is simply the number of LBOs in our sample each quarter. Scaled number of LBOs is the number of LBOs scaled by the number of public firms in the CRSP NYSE/AMEX/NASDAQ/ARCA index the prior quarter. RP IS, NDS *cay** is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1954Q1-2010Q3. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1952Q1-2013Q4. RP OOS, NDS *cay* is the risk premium estimated out-of-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill. RP IS, PCE *cay*, is the risk premium estimated in-sample with the PCE measure of *cay*, the D/P ratio and the three-month T-bill. RP OOS, PCE *cay* is the risk premium estimated out-of-sample with the PCE measure of *cay*, the D/P ratio and the three-month T-bill. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from COMPUSTAT and the yield on the Merrill Lynch High Yield Master II Bond Index. The HY spread is the yield on the Merrill Lynch High Yield Master II Bond Index less the three-month T-bill rate. The GZ spread is a measure of excess bond premia as defined in Gilchrist and Zakrajšek (2012). Transaction value is the sum of LBO transaction values in billions of 2009 dollars each quarter. Asset value is the sum of Assets for the LBO-deals in our sample in billions of 2009 dollars each quarter. Enterprise Value is the sum of Enterprise Value for the LBO-deals in our sample in billions of 2009 dollars each quarter. Equity β is target betas in our U.S. LBO sample based on monthly returns with a two-year estimation window. Unlevered β is the unlevered Equity β.

Variable	Obs	Mean	Std. Dev.	Min	Max
By Quarter					
Number of LBOs	137	9.72	5.94	0	29
Scaled number of LBOs	137	13.84	8.5	0	41.05
RP IS, NDS *cay**	117	4.03	5.55	-9.48	14.39
RP IS, NDS *cay*	137	5.43	3.23	-3.09	11.34
RP IS, PCE *cay*	137	6.03	4.68	-5.24	16.04
RP OOS, NDS *cay*	137	2.9	7.28	-13.46	19.04
RP OOS, PCE *cay*	137	4.27	5.45	-10.61	13.08
EBITDA spread	79	-7.51	2.79	-17.61	-3.91
HY spread	79	7.14	2.91	2.52	18.7
GZ spread	136	2.02	.98	.78	7.59
Transaction value	137	9.74	19.01	0	139.25
Asset value	137	10.15	17.14	0	106.98
Enterprise value	137	11.5	21.52	0	138.4
Full Sample					
Unlevered β	827	.79	.52	-.18	2.11
Equity β	827	1.01	.68	-.49	4.71

In Figure 1, the number of U.S. LBOs each quarter along with our U.S. in-sample NDS *cay* risk premium are displayed. Visually, the relationship between the two seems strong. The risk premium has been especially high during periods of lower LBO activity. Figure 1 does, however, suggest that while the risk premium has been relatively low in recent years, LBO activity has not followed suit.

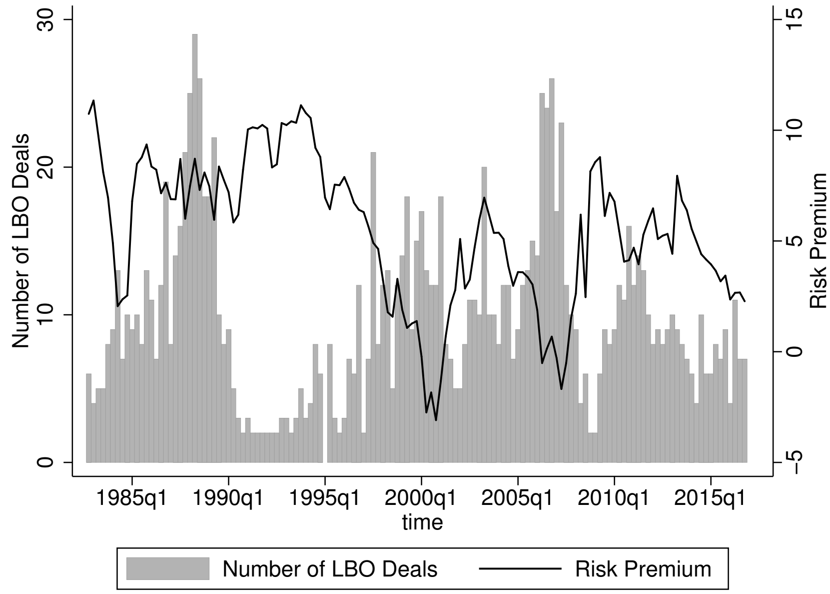

Figure 1. Number of U.S. public-to-private LBOs and the aggregate risk premium This figure shows the number of public-to-private U.S. LBOS between 1982Q4 and 2016Q4 along with an in-sample estimated aggregate risk premium based on NDS *cay*.

Figure 2 illustrates our scaled fundraising measure and the U.S. in-sample risk premium with NDS *cay* over time. Similar to the scaled number of LBOs measure, the scaled fundraising measure exhibits peaks in magnitude in the late 1980s and around 2007 although the 1980s peak is less pronounced compared to the scaled number of LBOs.

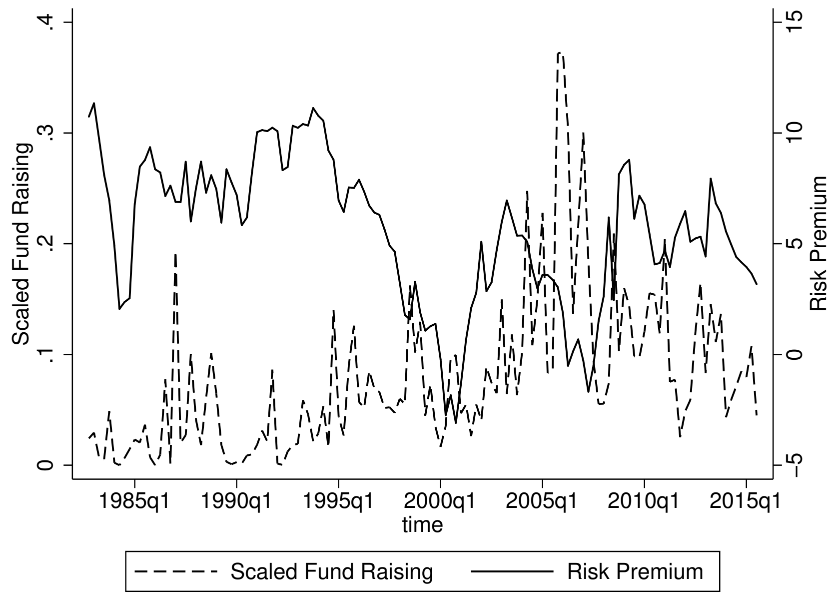

Figure 2. Scaled fundraising and the aggregate risk premium. This figure shows the sum of fundraising (fundraising assumed to occur 5 quarters prior to first investment in fund) scaled by the market value of the CRSP NYSE/AMEX/NASDAQ index in the U.S. between 1982Q4 and 2016Q4 along with an in-sample estimated aggregate risk premium based on NDS *cay*.

B. U.K. Sample Summary Statistics

Table II presents the equivalent summary statistics for our U.K. sample. This sample stretches from 1998Q2 to 2016Q4. For our U.K. sample there are 5.01 LBOs per quarter on average. The scaled activity measure shows that 0,205% of public U.K. firms on average are taken private each quarter. Like our U.S. sample, Transaction values, asset values and enterprise values are collected from the SDC Platinum database. The quarterly Transactions Values are on average \$1.69 bn. 15 of the 375 targets in the sample lack transaction value data. The quarterly asset values are on average \$2.73 bn. 14 of the 375 targets in the sample lack asset value data. The quarterly enterprise values are on average \$3.49 bn. 29 of the 375 targets in the sample lack enterprise value data. Similar to the U.S. all three measures have a large standard deviation compared to the mean.

Credit factors and our different U.K. risk premium estimations are also included in Table II. There are some notable differences in general between the out-of-sample and the in-sample risk premiums. The standard deviation between all risk measures is similar but there is a rather large difference in means and maximum values. The larger differences for U.K. compared to the U.S. could be due to the smaller risk premium estimation window before the regression window starts. The in-sample NDS and PCE *cay* risk premium estimations are very similar in both mean, minimum and maximum value. This is most likely due to that in the later part of the risk estimation window, *cay* has a rather low regression coefficient compared to the risk-free rate and the D/P ratio in the risk premium estimations.

Unlevered betas in our U.K. sample are on average 0.62 while the market betas are 0.77 for the 260 deals included in our regressions. The U.K. unlevered betas are on average 0.17 smaller than in our U.S. sample. The U.K. market equity betas are on average 0.24 smaller than in our U.S. sample. Potential bias considerations due to the relatively smaller sample size is in place as always but it is worth noting that the U.K. sample is more in line with what Frazzini and Pedersen (2014) find for their public-to-private LBO sample, i.e. market equity betas on average below 1.0.

Table II
Summary Statistics, U.K.

This table presents quarterly summary statistics for U.S. LBO activity, credit factors, accounting measures, transaction values and risk premiums. It also includes the Equity β and its corresponding Unlevered β. Number of LBOs is simply the number of LBOs in our sample each quarter. Scaled number of LBOs is the number of LBOs scaled by the number of public firms on the London Stock Exchange the prior quarter. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. RP OOS, NDS *cay* is the risk premium estimated out-of-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. RP IS, PCE *cay*, is the risk premium estimated in-sample with the PCE measure of *cay*, the D/P ratio and three-month U.K. government bond rate. RP OOS, PCE *cay* is the risk premium estimated out-of-sample with the PCE measure of *cay*, the D/P ratio and three-month U.K. government bond rate. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from the constituent list of the FTSE All Share Index from Datastream and the yield on the Merrill Lynch Euro High Yield Bond Index. The HY spread is the yield on the Merrill Lynch Euro High Yield Bond Index less the three-month U.K. government bond rate from Datastream. Transaction value is the sum of LBO transaction values in billions of 2009 pounds each quarter. Asset value is the sum of Assets for the LBO-deals in our sample in billions of 2009 pounds each quarter. Enterprise Value is the sum of Enterprise Value for the LBO-deals in our sample in billions of 2009 pounds each quarter. Equity β is target betas in our U.K. LBO sample based on monthly returns with a two-year estimation window. Unlevered β is the unlevered Equity β.

Variable	Obs	Mean	Std. Dev.	Min	Max
By Quarter					
Number of LBOs	75	5.01	3.62	0	19
Scaled number of LBOs	75	20.56	14.86	0	80.09
RP IS, NDS *cay*	75	5.47	7.82	-8.31	26.24
RP IS, PCE *cay*	75	5.37	7.63	-8.51	27.53
RP OOS, NDS *cay*	75	12.5	8.06	1.85	49.08
RP OOS, PCE *cay*	75	9.75	8.36	-3.35	43.33
EBITDA spread	75	-1.5	4.37	-16.28	3.58
HY spread	75	6.39	4.48	.7	23.17
Transaction value	75	1.69	2.99	0	17.58
Asset value	75	2.73	5.28	0	38.07
Enterprise value	75	3.49	9.59	0	76.51
Full Sample					
Unlevered β	260	.62	.45	-.38	1.93
Equity β	260	.77	.53	-.77	2.22

Figure 3 illustrates the number of U.K. LBOs per quarter in our sample along with the U.K. in-sample NDS *cay* risk premium. Similar to the U.S., the risk premium and LBO activity mirrors each other closely. Worth noting in contrast to Figure 1, the U.K. risk premium has been relatively high in recent years, aligning more closely to what the model of Haddad et al. (2017) predicts given the relatively low numbers of U.K. LBOs.

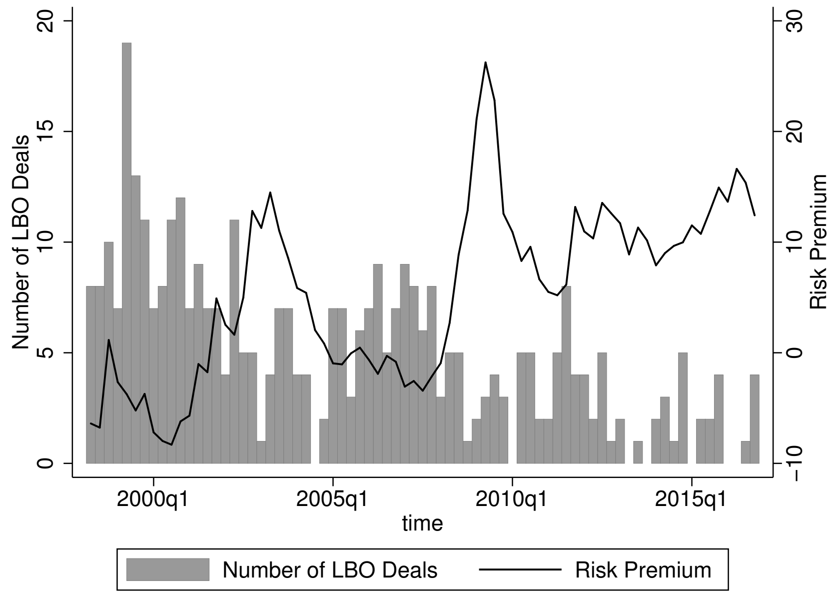

Figure 3. Number of U.K. public-to-private LBOs and the aggregate risk premium. This figure shows the number of public-to-private U.K. LBOS between 1998Q2 and 2016Q4 along with an in-sample estimated aggregate risk premium based on NDS *cay*.

IV. Results

A. U.S. LBO Activity Post-2011

We begin our analysis by focusing on our U.S. sample reexamining the results found by Haddad et al. (2017) in order to validate our sample selection criteria and the risk premium estimation. Table III shows that when using a risk premium estimated in-sample based on NDS *cay*, up until 2011Q4, there appears to be a clear relationship between LBO activity and the aggregate risk premium. Not only does the risk premium alone have a higher R-squared than all credit spreads regressed together, the risk premium is the only variable statistically significant when regressed together with the credit spreads (column (9)). The results correspond closely to the findings in Haddad et al. (2017). We label the risk premium estimated in-sample with NDS *cay* with an estimation window from 1954Q1 to 2010Q3, RP IS NDS *cay** to distinguish it from our identical but full sample version.

Table III
Drivers of U.S. LBO Activity, IS Risk Premium with NDS *cay**

This table presents coefficient estimates from regressions where the dependent variable is the number of U.S. LBOs scaled by the number of public firms in the CRSP NYSE/AMEX/NASDAQ/ARCA index the prior quarter and the independent variables are credit factors and an U.S. aggregate risk premium. RP IS, NDS *cay** is the risk premium estimated in-sample with the NDS measure of *cay* with an estimation window of 1954Q1-2010Q3. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from COMPUSTAT and the yield on the Merrill Lynch High Yield Master II Bond Index. The HY spread is the yield on the Merrill Lynch High Yield Master II Bond Index less the three-month T-bill rate. The GZ spread is a measure of excess bond premia as defined in Gilchrist and Zakrajšek (2012). In columns (1) and (7) we run the regression between 1982Q4 and 2011Q4. In columns (2) and (4) we run the regression between 1997Q2 and 2016Q4 where data is available for the HY spread and the EBITDA spread. In column (6) we run the regression between 1982Q4 and 2016Q3 where data is available for the GZ spread. Due to these limitations in data availability we run the regressions between 1997Q2 and 2011Q4 in columns (3) and (5), in column (8) we run the regression between 1997Q2 and 2016Q3 and in column (9) we run the regression between 1997Q2 and 2011Q4. Quarterly dummies to account for seasonality are included in each regression. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Scaled Number of Leveraged Buyouts

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
RP IS, NDS *cay**	-0.8359***		-0.7863**		-0.7950**		-0.9078***		-0.7655**
	(0.1945)		(0.3702)		(0.3698)		(0.1845)		(0.3635)
HY spread		-0.9871**	-0.5033**					-1.5591	-0.1423
		(0.4197)	(0.2264)					(1.0542)	(0.8203)
EBITDA spread				0.5409	0.7583***			-0.1634	0.7955
				(0.3758)	(0.2142)			(0.5102)	(0.5223)
GZ spread						-0.4411	-1.5808**	1.3806	0.4043
						(1.0291)	(0.6852)	(2.6409)	(3.0730)
Number of obs	117	79	59	79	59	136	117	78	59
R-squared	0.276	0.180	0.352	0.065	0.379	0.017	0.307	0.197	0.379

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Next, we test how well the risk premium estimated in-sample with NDS *cay*, with an estimation window from 1952Q1 to 2013Q4, explains LBO activity. In other words, we turn to our full sample extending beyond 2011. Table IV displays the results. There still appears to be a clear relationship between LBO activity and the aggregate risk premium. Again, the risk premium alone has a higher R-squared than all credit spreads regressed together and the risk premium is the only variable statistically significant when regressed together with the credit spreads (column (5)). However, the statistical significance of the risk premium regression coefficients is notably weaker, indicating that the relationship between the risk premium and LBO activity has been less pronounced in recent years.

Table IV
Drivers of U.S. LBO Activity, IS Risk Premium With NDS *cay*

This table presents coefficient estimates from regressions where the dependent variable is the number of U.S. LBOs scaled by the number of public firms in the CRSP NYSE/AMEX/NASDAQ/ARCA index the prior quarter and the independent variables are credit factors and an U.S. aggregate risk premium. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1952Q1-2013Q4. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from COMPUSTAT and the yield on the Merrill Lynch High Yield Master II Bond Index. The HY spread is the yield on the Merrill Lynch High Yield Master II Bond Index less the three-month T-bill rate. The GZ spread is a measure of excess bond premia as defined in Gilchrist and Zakrajšek (2012). In column (1) we run the regression between 1982Q4 and 2016Q4. In column (2) and (4) the regression window is limited to 1997Q2 to 2016Q4 where data on the Merrill Lynch High Yield Master II bond index and the EBITDA spread are available. In column (4) we run the regression between 1982Q4 and 2016Q3, the last quarter where data on the GZ spread is available. Column (5) includes all independent variables and is limited to 1997Q2 to 2016Q3 due to the earlier constraints. Quarterly dummies to account for seasonality are included in each regression. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Scaled Number of Leveraged Buyouts

	(1)	(2)	(3)	(4)	(5)
RP IS, NDS *cay*	-0.9370**	-0.7281*	-1.1242**	-1.0080***	-0.9995*
	(0.3833)	(0.3993)	(0.4703)	(0.3701)	(0.5146)
HY spread		-0.6513**			-0.6316
		(0.3129)			(0.9858)
EBITDA spread			0.5758**		0.6158
			(0.2857)		(0.6483)
GZ spread				-1.0198	1.5592
				(0.7222)	(2.1736)
Number of obs	137	79	79	136	78
R-squared	0.140	0.232	0.231	0.158	0.251

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Turning to the relationship between target betas and the risk premium and LBO activity, we start by regressing our U.S. target betas on the U.S. risk premium estimated in-sample based on NDS *cay* up-until 2011. Unsurprisingly, as Table V shows, a lower risk premium is associated with higher target betas as found by Haddad et al. (2017) and the result is significant at the 1% level. As also found by Haddad et al. (2017), higher LBO activity is associated with higher betas. This result is significant at the 1% level both during the 1982Q4 - 2011Q4 period (column (1)) and the 1982Q4 - 2016Q4 period (column (2)).

We then turn to our full sample NDS *cay* risk premium estimated in-sample (Table V, column (4)). The direction of the regression coefficient is the same as for the shorter estimation and regression window, but the result is statistically insignificant suggesting at a first glance that target betas in the latest part of our sample period does not behave quite the same way as what they did leading up to 2011. However, given the already noted weaker link between the risk premium and LBO activity, this should come as no surprise. The finding that LBO target betas are higher when LBO activity is high still holds true. In sum, it thus seems like the risk premium has been a weaker determinant of LBO activity in recent years while LBO activity itself has continued to influence the magnitude of target betas. Figure 4 in Appendix A displays the Kernel density of our U.S. target betas for the top and bottom quartile in-sample NDS *cay* risk premium periods. It is difficult to discern a clear difference in relative mass above 1 for the high and low risk premium quartiles when including targets post-2011.

Table V
Drivers of U.S. LBO Target Betas

This table presents coefficient estimates from regressions where the dependent variable is the unlevered target betas in our U.S. LBO sample based on monthly returns with a two-year estimation window and the independent variables are an activity measure and U.S. aggregate risk premiums. Scaled activity is the number of U.S. LBO deals each quarter scaled by the number of public U.S. firms. RP IS, NDS *cay** is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1954Q1-2010Q3. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1952Q1-2013Q4. In columns (1), and (3) we run the regression for firms that was acquired between 1982Q4 and 2011Q4. In column (2), and (4) we run the regression in the full sample, i.e. for firms acquired between 1982Q4 and 2016Q4. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Unlevered Target Betas

	(1)	(2)	(3)	(4)
Scaled activity	0.0100***	0.0086***		
	(0.0021)	(0.0021)		
RP IS, NDS *cay**			-0.0105***	
			(0.0041)	
RP IS, NDS *cay*				-0.0028
				(0.0069)
Number of obs	716	827	716	827
R-squared	0.031	0.023	0.012	0.000

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

B. U.K. LBO Activity

In Table VI, the results from testing the relationship between LBO activity and risk premium in the U.K are displayed. The risk premium used is a U.K risk premium estimated in-sample with NDS *cay*. A higher risk premium is associated with lower activity and the result is robust to the inclusion of the U.K. High Yield Spread and U.K. EBITDA spread (Column (7)). In each specification, all the coefficients for the risk premium are significant at the 1% level. Neither the High Yield Spread nor the EBITDA spread are statistically significant on their own. The regression based solely on the risk premium has an R-squared that is significantly higher than the R-squared for the regression including only the two credit spreads. Furthermore, the R-squared is significantly higher than what we find for the U.S. when the risk premium is based on NDS *cay* and estimated in-sample. The results found in the U.S. sample thus holds true in the U.K. as well, using an in-sample risk premium measure with our estimated U.K. NDS *cay*. As noted in Section III, the relationship between the risk premium and LBO activity in both the U.S. and the U.K. market appears strong only by looking at Figure 1 and 3.

Table VI
Drivers of U.K. LBO Activity, IS Risk Premium With NDS *cay*

This table presents coefficient estimates from regressions where the dependent variable is the number of U.K. LBOs scaled by the number of public firms on the London Stock Exchange the prior quarter and the independent variables are credit factors and an aggregate risk premium. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from the constituent list of the FTSE All Share Index from Datastream and the yield on the Merrill Lynch Euro High Yield Bond Index. The HY spread is the yield on the Merrill Lynch Euro High Yield Bond Index less the three-month U.K. government bond rate from Datastream. We run all regression between 1998Q2 and 2016Q4. Quarterly dummies to account for seasonality are included in each regression. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Scaled Number of Leveraged Buyouts

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
RP IS, NDS *cay*	-1.2098***		-1.5083***		-1.2701***		-1.7634***
	(0.2450)		(0.2102)		(0.2008)		(0.2847)
HY spread		-0.1142	1.1171***			-1.6349**	2.5379***
		(0.3894)	(0.2424)			(0.7183)	(0.9720)
EBITDA spread				-0.3372	-0.6922***	-1.7929*	1.4296
				(0.5557)	(0.1855)	(0.9416)	(0.8954)
Number of obs	75	75	75	75	75	75	75
R-squared	0.449	0.047	0.537	0.056	0.490	0.115	0.566

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

As Figure 3 shows, when basing the activity measure on the scaled number of deals, there is a clear peak in activity in 1999/2000. As mentioned in Section I.A, Renneboog and Vansteenkiste (2017) highlight the difference between the U.K. volume of public-to-private LBOs in terms of number and value. The late 1990s was characterized by deals where small targets were acquired. In order to assess if the results differ if a value measure of activity is used instead, we construct a measure of activity defined as the log of the sum of LBO transaction values each quarter. As Table VII shows, when this measure of activity is used the regression coefficients remains negative for the risk premium and the significance levels are intact, although the R-squared when regressing this alternative activity measure solely on the risk premium drops in magnitude. In contrast, the R-squared from the regression where only the High Yield Spread and the EBITDA spread are used increase, approaching the size of the R-squared in the regression including only the risk premium. However, the coefficients of the two credit factors remain statistically insignificant on their own.

Table VII
Drivers of U.K. LBO Activity, IS Risk Premium With NDS *cay*

This table presents coefficient estimates from regressions where the dependent variable is the log of the sum of U.K. LBO transaction values each quarter and the independent variables are credit factors and an U.K. aggregate risk premium. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from the constituent list of the FTSE All Share Index from Datastream and the yield on the Merrill Lynch Euro High Yield Bond Index. The HY spread is the yield on the Merrill Lynch Euro High Yield Bond Index less the three-month U.K. government bond rate from Datastream. We run all regression between 1998Q2 and 2016Q4. Quarterly dummies to account for seasonality are included in each regression. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Log Transaction Values

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
RP IS, NDS *cay*	-0.1630***		-0.1783***		-0.1694***		-0.1509***
	(0.0355)		(0.0354)		(0.0313)		(0.0342)
HY spread		-0.0883	0.0573			-0.4526***	-0.0956
		(0.0587)	(0.0571)			(0.1189)	(0.1391)
EBITDA spread				-0.0265	-0.0739	-0.4295***	-0.1538
				(0.0724)	(0.0490)	(0.1365)	(0.1230)
Number of obs	75	75	75	75	75	75	75
R-squared	0.299	0.068	0.307	0.046	0.315	0.204	0.319

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

We also examine the effect of the risk premium and LBO activity on target betas for the U.K. Table VIII displays the results. The risk premium estimated in-sample with NDS *cay* is statistically significant at the 10% level although the sign of the coefficient is reversed compared to the U.S. implying that a higher risk premium is associated with higher betas. Table VIII also shows the results from regressing target betas on LBO activity. Again, the sign of the coefficient is reversed compared to the U.S. While the result is statistically significant at the 1% level, it might not paint the full picture. When instead regressing the transaction value measure of LBO activity on the target betas, the sign is positive yet statistically insignificant.

In Table XV in Appendix A the same regression is performed on weekly betas. The results are largely the same. The regression coefficients for scaled activity (column (1)) and log transaction values (column (2)) have the same sign and are statistically significant at the same levels as for monthly betas. The regression coefficient for the risk premium estimated in sample using NDS *cay* (column (3)) has the same sign but is not statistically significant. Figure 5 in Appendix A displays the kernel density of our U.K. target betas for the top and bottom quartile in-sample NDS *cay* risk premium periods. In contrast to the U.S. it seems like relatively more mass is concentrated at higher beta levels for the high risk premium quartile compared to the low quartile risk premium. In Section V, we give these conflicting results some more attention and discuss our findings in contrast to the results found in the U.S.

Table VIII
Drivers of U.K. LBO Target Betas

This table presents coefficient estimates from regressions where the dependent variable is the unlevered target betas in our U.K. LBO sample based on monthly returns with a two-year estimation window and the independent variables are two activity measures and an U.K. aggregate risk premium. Scaled activity is the number of U.K. LBO deals each quarter scaled by the number of public U.K. firms. Log transaction value is the log of the sum of transaction value of all LBO deals in our U.K sample each quarter. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. Newey-West standard errors (4 lags) in parentheses.

	Dep. Var.: Unlevered Target Betas		
	(1)	(2)	(3)
Scaled activity	-0.0057***		
	(0.0012)		
Log transaction value		0.0222	
		(0.0196)	
RP IS, NDS *cay*			0.0055*
			(0.0033)
Number of obs	260	260	260
R-squared	0.051	0.004	0.009

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

C. Alternative Risk Premiums

To test if the results for the U.S. and U.K. samples change when using alternative methods for estimating the risk premium, we run our regressions with risk premiums estimated out-of-sample and by using PCE instead of NDS *cay*. As mentioned in Section II.C, stating which method that is most appropriate would be highly subjective given the conflicting views in existing literature regarding return forecasting. Therefore, alternatives to the base case, i.e. the risk premium estimated in-sample with NDS *cay*, deserves some attention.

In Table IX, the results when regressing LBO activity on alternative aggregate risk premium measures for the U.S. sample are displayed. The PCE risk premiums do not show statistical significance when regressed in isolation. The out-of-sample NDS risk premium is statistically significant at the 5% level when regressed in isolation and remains significant when including the credit measures in the regression.

Table IX
Drivers of U.S. LBO Activity, Alternative Risk Premiums

This table presents coefficient estimates from regressions where the dependent variable is the number of U.S. LBOs scaled by the number of public firms in the CRSP NYSE/AMEX/NASDAQ/ARCA index the prior quarter and the independent variables are credit factors and measures of the U.S. aggregate risk premiums. RP OOS, NDS *cay* is the risk premium estimated out-of-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill. RP IS, PCE *cay*, is the risk premium estimated in-sample with the PCE measure of *cay*, the D/P ratio and the three-month T-bill. RP OOS, PCE *cay* is the risk premium estimated out-of-sample with the PCE measure of *cay*, the D/P ratio and the three-month T-bill. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from COMPUSTAT and the yield on the Merrill Lynch High Yield Master II Bond Index. The HY spread is the yield on the Merrill Lynch High Yield Master II Bond Index less the three-month T-bill rate. The GZ spread is a measure of excess bond premia as defined in Gilchrist and Zakrajšek (2012). In column (1), (3) and (5) we run the regression between 1982Q4 and 2016Q4. In column (2), (4) and (6) we run the regression between 1997Q2 and 2016Q4 where data on the Merrill Lynch High Yield Master II bond index and the EBITDA spread are available. Quarterly dummies to account for seasonality are included in each regression. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Scaled Number of Leveraged Buyouts

	(1)	(2)	(3)	(4)	(5)	(6)
RP IS, PCE *cay*	-0.3508	-0.0685				
	(0.3053)	(0.2931)				
EBITDA spread		-0.1234		0.4956		-0.1300
		(0.5773)		(0.3963)		(0.4749)
HY spread		-1.4617		-1.0853		-1.0017
		(1.0421)		(0.8894)		(0.8244)
GZ spread		1.3670		2.7563		1.2885
		(2.5835)		(2.4194)		(2.4282)
RP OOS, NDS *cay*			-0.3675**	-0.5679***		
			(0.1674)	(0.1955)		
RP OOS, PCE *cay*					-0.1413	-0.3705
					(0.3369)	(0.2879)
Number of obs	137	78	137	78	137	78
R-squared	0.051	0.198	0.113	0.340	0.022	0.239

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table X displays the corresponding regression results for our U.K. sample. Interestingly, both the signs of the risk premium coefficients and their statistical significance at the 1% level remains intact. We leave a more thorough analysis of these results to Section V but it is worth noting that the comparability of the U.K. and U.S. results, apart from being affected by different sample sizes, likely will be affected by differences in how *cay* is constructed.

Table X
Drivers of U.K. LBO Activity, Alternative Risk Premiums

This table presents coefficient estimates from regressions where the dependent variable is the number of U.K. LBOs scaled by the number of public firms on the London Stock Exchange the prior quarter and the independent variables are credit factors and measures of the U.K. aggregate risk premiums. RP OOS, NDS *cay* is the risk premium estimated out-of-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. RP IS, PCE *cay*, is the risk premium estimated in-sample with the PCE measure of *cay*, the D/P ratio and three-month U.K. government bond rate. RP OOS, PCE *cay* is the risk premium estimated out-of-sample with the PCE measure of *cay*, the D/P ratio and three-month U.K. government bond rate. The EBITDA spread is the difference between the median public firm's EBITDA/EV ratio from the constituent list of the FTSE All Share Index from Datastream and the yield on the Merrill Lynch Euro High Yield Bond Index. The HY spread is the yield on the Merrill Lynch Euro High Yield Bond Index less the three-month U.K. government bond rate from Datastream. We run all regression between 1998Q2 and 2016Q4. Quarterly dummies to account for seasonality are included in each regression. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Scaled Number of Leveraged Buyouts

	(1)	(2)	(3)	(4)	(5)	(6)
RP IS, PCE *cay*	-1.2511***	-1.7587***				
	(0.2597)	(0.2786)				
HY spread		2.2906**		-0.3910		-0.9402
		(0.9213)		(0.7681)		(0.7474)
EBITDA spread		1.2032		-1.2579		-1.5892*
		(0.8464)		(0.8529)		(0.8601)
RP OOS, NDS *cay*			-0.5397***	-0.7043***		
			(0.1568)	(0.1766)		
RP OOS, PCE *cay*					-0.4263***	-0.5040**
					(0.1524)	(0.2139)
Number of obs	75	75	75	75	75	75
R-squared	0.457	0.565	0.132	0.198	0.104	0.170

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

In Table XI we turn our focus to the relationship between target betas and our alternative risk premiums. Columns (1) - (3) displays the results for the U.K. sample while column (4) - (6) displays the U.S. results. Starting with the U.S., the sign of the coefficients remains negative and not statistically significant for all risk premium specifications. The out-of-sample risk premiums performs slightly better than their in-sample PCE *cay* counterpart but show no statistical significance.

For the U.K., the in-sample PCE *cay* risk premium performs similar to how the in-sample NDS *cay* risk premium did. The sign of the regression coefficient is reversed compared to the U.S. implying that target betas are higher when the risk premium is high. The result is significant at the 5% level. The results for the out-of-sample risk premiums are statistically insignificant.

Table XI
Drivers of Betas (U.S. and U.K.), Alternative Risk Premiums

This table presents coefficient estimates from regressions where the dependent variable is the unlevered target betas in our LBO samples based on monthly returns with a two-year estimation window and the independent variables are aggregate risk premiums. In columns (1), (2) and (3) we run the regression for our U.K. sample. In columns (4), (5) and (6) we run the regression for our U.S. sample. RP OOS, NDS *cay* is the risk premium estimated out-of-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate (three-month T-bill for our U.S. sample). RP IS, PCE *cay*, is the risk premium estimated in-sample with the PCE measure of *cay*, the D/P ratio and three-month U.K. government bond rate (three-month T-bill for our U.S. sample). RP OOS, PCE *cay* is the risk premium estimated out-of-sample with the PCE measure of *cay*, the D/P ratio and three-month U.K. government bond rate (three-month T-bill for our U.S. sample). Newey-West standard errors (4 lags) in parentheses.

	Dep. Var.: Unlevered Target Betas					
	(1)	(2)	(3)	(4)	(5)	(6)
RP IS, PCE *cay*	0.0067**			-0.0019		
	(0.0031)			(0.0041)		
RP OOS, NDS *cay*		0.0029			-0.0038	
		(0.0031)			(0.0027)	
RP OOS, PCE *cay*			-0.0008			-0.0046
			(0.0033)			(0.0032)
Number of obs	260	260	260	827	827	827
R-squared	0.012	0.002	0.000	0.000	0.003	0.003

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

D. *Betas over time*

Additionally, we test how the average beta of LBO targets has developed over time. In Table XII the results for our U.S. sample are displayed and in Table XIII the results for our U.K. sample are displayed. The variable *Time* increases one unit per quarter.

For the U.S. sample, *Time* is statistically significant at the 1% level for all columns except for the regression including the in-sample NDS *cay** risk premium (Table XII column (3)). In all regressions the direction of *Time* is positive which implies that betas on average have increased over time. It is also worth noting that controlling for time removes the statistical significance for the in-sample NDS *cay** risk premium.

Table XII
Drivers of U.S. Betas, Controlling for Time

This table presents coefficient estimates from regressions where the dependent variable is the unlevered target betas in our U.S. LBO sample based on monthly returns with a two-year estimation window and the independent variables are an activity measure, a time-variable and U.S. aggregate risk premiums. Scaled activity is the number of U.S. LBO deals each quarter scaled by the number of public U.S. firms. *Time* is a variable that increases with 1 for each quarter. RP IS, NDS *cay** is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1954Q1-2010Q3. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1952Q1-2013Q4. In columns (1), and (3) we run the regression for all firms acquired between 1982Q4 and 2011Q4. In column (2), (4) and (5) we run the regression in the full sample, i.e. for all firms acquired between 1982Q4 and 2016Q4. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Unlevered Target Betas

	(1)	(2)	(3)	(4)	(5)
Scaled activity	0.0102***	0.0100***			
	(0.0020)	(0.0019)			
Time	0.0015***	0.0016***	0.0008	0.0013***	0.0012***
	(0.0005)	(0.0004)	(0.0008)	(0.0005)	(0.0004)
RP IS, NDS *cay**			-0.0069		
			(0.0059)		
RP IS, NDS *cay*				0.0043	
				(0.0073)	
Number of obs	715	827	715	827	827
R-squared	0.042	0.037	0.013	0.009	0.008

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

For the U.K. sample, *Time* is statistically significant at the 5% level for columns (3) and (4), statistically significant at the 1% level together with log transaction value (column (2)) and lacks statistical significance for the regression including LBO activity (column (1)). In all regressions the direction of time is positive which implies that betas on average has increased over time. In sum, it seems like there is a relationship with time and beta in the U.K. sample as well.

Table XIII
Drivers of U.K. LBO Target Betas, Controlling for Time

This table presents coefficient estimates from regressions where the dependent variable is the unlevered target betas in our U.K. LBO sample based on monthly returns with a two-year estimation window, a time-variable and the independent variables are two activity measures and an U.K. aggregate risk premium. Scaled activity is the number of U.K. LBO deals each quarter scaled by the number of public U.K. firms. *Time* is a variable that increases with 1 for each quarter. Log transaction value is the log of the sum of transaction value of all LBO deals in our U.K sample each quarter. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. We run all regressions for all firms acquired between 1998Q2 and 2016Q4. Newey-West standard errors (4 lags) in parentheses.

	Dep. Var.: Unlevered Target Betas			
	(1)	(2)	(3)	(4)
Scaled activity	-0.0055***			
	(0.0012)			
Time	0.0028*	0.0039**	0.0034*	0.0032**
	(0.0016)	(0.0017)	(0.0018)	(0.0016)
Log transaction value		0.0380*		
		(0.0199)		
RP IS, NDS *cay*			-0.0007	
			(0.0036)	
Number of obs	260	260	260	260
R-squared	0.066	0.032	0.020	0.020

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

E. Fundraising and the Equity Risk Premium

In Table XIV we regress our fundraising measure on our different aggregate risk premium estimations. Since our data does not allow us to determine the timing of the actual investment decision by the limited partners in relation to the first investment by the fund, we use a number of lags on the risk premium to capture a variety of alternative timings. Column (1) displays the results when using no lag on the risk premium. Each column thereafter adds a one quarter lag. Apart from the risk premium estimated out-of-sample with PCE *cay*, the result for all regressions is statistically significant if applying anything in-between zero and eight lags. The R-Squared is highest in each regression if five lags are applied apart from the out-of-sample NDS *cay* risk premium where the highest R-Squared is achieved if using six lags. Again, important to note is that the actual time of fundraising is unknown in our data but it would be reasonable to assume that the first investment does not take place the same quarter as the fund is closed on average. An increasing R-Squared as the lags increase would thus make sense if the private equity funds require some time before finding and finalizing its first investment. At the same time, one would expect that the time between a fund's closing and its first investment would not be several years given the limited life span of a typical fund. The peak in R-squared at the five to six lag interval would indicate that if the risk premium in fact plays a major role in the decision by limited partners to invest in private equity, that decision might on average be made roughly one and a half years before the fund's first investment. It should be stressed that there are several potential biases in these results which we will cover in more detail in Section V.

Table XIV
U.S. PE Fundraising and the Equity Risk Premium

This table presents coefficient estimates from regressions where the dependent variable is the sum of the funds raised in each quarter (proxied by using first investment date and applying a lag) scaled by the Market Value recorded in the CRSP NYSE/AMAX/NASDAQ index and the independent variables are aggregate U.S. aggregate risk premiums. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill with an estimation window of 1952Q1-2013Q4. RP OOS, NDS *cay* is the risk premium estimated out-of-sample with the NDS measure of *cay*, the D/P ratio and the three-month T-bill. RP IS, PCE *cay*, is the risk premium estimated in-sample with the PCE measure of *cay*, the D/P ratio and the three-month T-bill. RP OOS, PCE *cay* is the risk premium estimated out-of-sample with the PCE measure of *cay*, the D/P ratio and the three-month T-bill. In column (1) no lag is applied. In column (2) funds raised are moved as if the fundraising occurred one quarter earlier. Column (3)-(9) adds one further lag for each column. For example, in column (9) funds raised are moved as if the fundraising occurred eight quarters earlier. Quarterly dummies to account for seasonality are included in each regression. Newey-West standard errors (4 lags) in parentheses. R-Squared in brackets.

Dep. Var.: Scaled Number of Fundraising

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
RP IS, NDS *cay*	-0.0091***	-0.0089**	-0.0088**	-0.0092**	-0.0093**	-0.0085**	-0.0067**	-0.0064**	-0.0054
	(0.0032)	(0.0035)	(0.0036)	(0.0038)	(0.0038)	(0.0035)	(0.0032)	(0.0032)	(0.0033)
	[0.180]	[0.177]	[0.170]	[0.181]	[0.185]	[0.155]	[0.101]	[0.089]	[0.062]
RP IS,PCE *cay*	-0.0058***	-0.0062***	-0.0060**	-0.0064**	-0.0069**	-0.0064**	-0.0047*	-0.0044*	-0.0034
	(0.0022)	(0.0023)	(0.0025)	(0.0027)	(0.0028)	(0.0027)	(0.0025)	(0.0025)	(0.0027)
	[0.156]	[0.176]	[0.161]	[0.173]	[0.194]	[0.163]	[0.092]	[0.077]	[0.047]
RP OOS, NDS *cay*	-0.0046***	-0.0047***	-0.0048***	-0.0051***	-0.0054***	-0.0056***	-0.0053***	-0.0054***	-0.0052***
	(0.0012)	(0.0012)	(0.0012)	(0.0013)	(0.0014)	(0.0013)	(0.0013)	(0.0014)	(0.0016)
	[0.233]	[0.246]	[0.256]	[0.287]	[0.322]	[0.340]	[0.298]	[0.296]	[0.260]
RP OOS, PCE *cay*	-0.0051*	-0.0054**	-0.0047	-0.0055*	-0.0062**	-0.0058**	-0.0042	-0.0040	-0.0035
	(0.0027)	(0.0027)	(0.0029)	(0.0031)	(0.0031)	(0.0029)	(0.0029)	(0.0029)	(0.0031)
	[0.111]	[0.120]	[0.095]	[0.118]	[0.144]	[0.126]	[0.070]	[0.064]	[0.047]
Number of obs	137	136	135	134	133	132	131	130	129

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

V. Discussion of Results and Potential Biases

Before we discuss our findings in greater detail we will present a brief summary of our findings. To begin with, our results for the U.S. market suggest that while the relationship between LBO activity and the aggregate risk premium is still evident in the data, the link after 2011 seems to be weaker than what it has been previous years (see Table III and IV). A weaker link post-2011 is also evident when regressing target betas on the risk premium. However, the relationship between target betas and LBO activity remains strong (see Table V). Using PCE *cay* in the risk premium estimation seems to have a strong impact on the statistical significance of the results, although the sign of the regression coefficients remains intact. The out-of-sample risk premium with NDS *cay* yields results that are in line with its in-sample counterpart when regressed on LBO activity (see Table IX and XI). Overall, our tests on the U.K. market support the U.S. results. A higher risk premium is associated with lower LBO activity regardless of which activity measure is used (see Table VI and X). However, when U.K. target betas are regressed on LBO activity or the U.K. risk premiums, the results indicate that the opposite relationship to what is found in the U.S. holds true (see Table VIII). Estimating the betas on a weekly, instead of a monthly, basis does not affect the results (see Table XV in Appendix A). The alternative U.K. risk premiums do not seem to change the overall results either (see Table XI). On both markets, controlling the beta regressions with time indicates that target betas have increased on average over our sample periods (see Table XII and XIII). Finally, our fundraising measure on the U.S. LBO market indicates that there also is a potentially significant relationship between the aggregate risk premium and fundraising (see Table XIV).

The slightly differing results when estimating the risk premium out-of-sample or in-sample deserve some further discussion. First of all, it is arguably impossible to establish which risk premium estimation method objectively best capture the beliefs of future market returns at any point in time. If returns are not forecastable using the dividend-price ratio, *cay* and risk-free rate, and practitioners therefore do not rely on them, our out-of-sample risk premium measure would not accurately reflect market expectations. On the other hand, an estimation method colored by look-ahead bias might not be any stronger if the aim is to accurately proxy for market expectations ex-ante the buyout decision. However, our regression results indicate that the choice between using an in-sample or out-of-sample risk premium has a smaller effect on the outcome than the choice between using NDS or PCE *cay*.

When it comes to the differing results when using PCE versus NDS expenditure to estimate *cay*, it is worth noting once more that a definite answer to which measure is to be preferred is still absent in existing literature. As mentioned in Section II.C, consumption is unobservable but there is no indication that a PCE expenditure-based *cay* in theory is an inferior forecaster of future returns. The measures do however differ more than one might expect. The two measures of U.S. *cay* are far from aligning perfectly when compared to each other, as shown in Figure 6 in Appendix A. In some periods, the PCE measure of *cay* can be negative while the NDS measure is positive and vice versa. Lettau and Ludvigson (2015) demonstrate how the ratio of NDS over PCE has changed since 1952, moving from around 90% in 1952 to around 75% in 2014. With the ratio changing over time, it is not surprising that there is not a perfect relationship between the resulting *cay* measures. Given the relationship between the two measures and the high weight on *cay* in the U.S. risk premium regressions, it is rather expected that the NDS *cay* risk premium and the PCE *cay* risk premium differ, resulting in somewhat different regression coefficients. For the U.K., the difference between the risk premium based on NDS *cay* and the risk premium based on PCE *cay* is much smaller than for the U.S. Mainly, this is due to the low coefficient of *cay* in the risk premium estimation. However, as Figure 6 in Appendix A illustrates, even if this was not the case, the difference between the two risk premium measures would likely be smaller for the U.K. than the U.S. given the smaller spread between NDS *cay* and PCE *cay*. Most likely, this depends on a shorter estimation window for the U.K. *cay*s than for the U.S. *cay*s since the difference over time in the ratio of NDS over PCE should not be as clear in the U.K. Although identifying what separates NDS and PCE consumption helps explain the differing regression results, no attempt is made in this paper to distinguish which *cay* measure most accurately reflects the consumption-wealth ratio. Thus, we leave it to future research to potentially discern their relative strength in forecasting future market returns.

Regarding the less pronounced relationship between the equity risk premium and LBO activity found in our U.S. sample post-2011, it is possible that *cay*, or any of the other forecasting regressors, has been a weak forecaster of future returns in recent years and thus a poor indicator of the equity risk premium. Barrell et al. (2015) find that Italian and U.K. consumption responded differently to the financial crisis of 2008 which dramatically affected financial wealth. Consumption behavior in the U.K. were more sticky, responding slowly to the wealth decrease. A similar effect might be found in U.S. data. For the U.K., around the time of the 2008 financial crisis, our estimated *cay* loses statistical significance in its predictive ability on market excess returns. Regardless of the underlying reasons, the U.S. data indicates a very low risk premium in recent years, almost as low as just before the financial crisis of 2008, as shown in Figure 1. However, the number of LBO deals have been modest. In contrast, Figure 3, providing an overview of booms and busts in the U.K. LBO market, indicates that the risk premium has been higher than the historical average post-2011, more accurately reflecting the lower LBO activity also apparent in the U.K. LBO market. This discrepancy raises the question why the U.K. and U.S. risk premiums have behaved differently in recent years.

A potential source for the differing results is discrepancies in the strength of factors used to forecast returns. For instance, when comparing the dividend-price ratio's ability to forecast returns across countries, Campbell (2003) finds that the results differ substantially. In the U.S., Australia and particularly in the U.K., the dividend-price ratio is a strong predictor of future returns. However, limited forecasting capabilities are found in France, Germany and Japan. The data used in Campbell (2003) ends in 1999. Cornell (2014) revisits Campbell's findings and arrives at a similar conclusion. The dividend-price ratio has a particularly strong forecasting capability in the U.K. and the U.S. In his regressions, the R-squared is roughly three times higher in the U.K. compared to the U.S. Furthermore, the author finds that the dividend-price ratio in the U.K. and the U.S. is a weak predictor of dividend growth which should imply that it, at least in theory, is a suitable factor for forecasting returns[14]. In our data we also find that the dividend-price ratio contributes significantly more to the R-squared in our U.K. risk premium estimations compared to the U.S. where *cay* is more dominant. In recent years there has been an exceptionally low level of the risk-free rate and decreasing values of *cay*. As a result of this, a risk premium estimate where the dividend-price ratio, rather than *cay*, dominates the result will lead to a higher risk premium that

[14] From the approximate present value identity presented in Campbell and Shiller (1988), the dividend-price ratio could either predict future returns, future dividend growth or some combination of both.

more accurately reflects recent years' buyout volume. We cannot escape from the possibility that our results are affected by the lack of available historical data on the inputs used to construct the U.K. *cay* measures. Our estimated *cay*s most likely lack the robustness of its U.S. counterparts.

Regarding the conflicting results found when regressing U.K. target betas on the U.K. scaled number of LBOs activity measure, there are some aspects to consider. Renneboog and Vansteenkiste (2017) state that anecdotal evidence supports that the peak in number of public-to-private transactions in the late 1990s was a result of temporarily undervaluation. Smaller firms in particular, discarded by institutional investors, opted for private equity. While the model of Haddad et al. (2017) is agnostic regarding deal size, making the authors favor the scaled activity measure, the value of deals and number of deals follow each other much closer in the U.S. than in the U.K. It lies beyond the scope of this paper to offer any guidance regarding which measure of activity is to be preferred. However, the unmatched peak in number of U.K. transactions in the late 1990s should perhaps be considered an outlier. Concluding that betas follow an opposite pattern compared to the U.S., being on average higher when LBO activity is low, would thus be difficult based solely on our data. The results for the scaled activity measure are nonetheless coherent with the results found when regressing U.K. betas on the U.K. in-sample NDS *cay* risk premium, although the statistical significance of the results is limited to the 10% level. The alternative U.K. risk premiums paint the same picture overall and the reversed direction compared to the U.S. remains if the betas are constructed using weekly return data. In sum, the relationship between target betas and LBO activity might in fact behave differently in the U.S. compared to the U.K. One could also argue that the lower average U.K. target betas are more consistent in size with the LBO target betas reported by Frazzini and Pedersen (2014). Although, as mentioned in Section I.B, the authors do not explicitly claim that private equity firms engage in betting against beta strategies, it might be the case that the U.K. private equity firms consider betas and their role in contributing to private equity returns differently than their U.S. counterparts.

Apart from the 1990s, the two countries nonetheless display similar patterns in number of transactions, and it is possible that any large deviations will become smaller looking forward. In fact, anecdotal evidence indicates an increasingly integrated global private equity market. For instance, looking at the 2017 annual report of well-established limited partner the Canadian Pension Plan Investment Board (CPPIB), it is clear that their portfolio diversification has increased over time. In 2016, global assets constituted 83.5% of their total investment portfolio. The corresponding figure in the year 2000 was 18.3% (CPPIB, 2017). Of course, CPPIB might not be a representative example of the average limited partner. However, according to the 2016 European Private Equity Activity report by Invest Europe, a trade association surveying private equity activity in Europe, 36.2% of the capital raised by buyout funds in 2016 came from North America. In total, 52.2% of the capital raised by buyout funds came from outside of Europe (Invest Europe, 2017)[15]. Thus, one could even speculate that a local U.K. risk premium does not best capture the risk premium

[15] Invest Europe, formerly European Private Equity & Venture Capital Association, claims to cover 88% of the 600 billion capital under management by private equity funds in Europe.

relevant for an average buyout investor going forward.

For both our samples we find that the average beta has increased over time. As mentioned in Section I.A, Strömberg (2008) notes that there seems to have been a shift in the types of industries private equity firms invest in. The author offers two possible explanations to this shift. It could either reflect a change in industry mix for the economy as a whole, or it could reflect a deliberate shift by private equity firms away from traditional private equity targets. Aldatmaz and Brown (2018) find that after private equity investments, changes occur within the target's country and industry. These changes consist of increases in capital expenditures, labor productivity, employment and profitability for publicly listed companies. They suggest that after a private equity investment is made, positive externalities arise and are absorbed by companies within the same industry. Bernstein et al. (2017) find similar dynamics. Total productivity grows more quickly in an industry after private equity investments in the same industry. Furthermore, these industries are found to be less exposed to aggregate shocks.

The buyout decision model (see Equation 1) constructed by Haddad et al. (2017) includes a factor, p_H, capturing the output increase of a firm if a private equity fund were to acquire it. If one assumes that there is a certain set of techniques a private equity firm apply to improve a firms output, some interesting potential interpretations arise. If there are positive externalities within an industry after a private equity investment occurs, p_H for the firms within the same industry is likely to decrease over time since the set of the performance improvements that the private equity firm would implement already has started to materialize. A decrease in p_H implies a lower deal surplus and thus deal likelihood through the performance channel. The resulting effects from a decreasing p_H throughout the industry would then lead to a decreasing likelihood of a buyout for firms within the entire industry. This offers a potential explanation to the rising betas and shifting industry mix. If p_H has decreased in traditional private equity industries, buyout deals in those industries have become less attractive over time. A shift from the traditional private equity industries could therefore have taken place. If private equity firms have traditionally favored low beta targets, the decreasing performance gains attainable in low beta industries might have forced private equity firms to seek investment opportunities in higher beta industries. Such a development would be consistent with the findings in our data. However, it should be stressed that this is only one of several possible interpretations and a hypothesis our data prevents us from testing empirically.

Finally, our fundraising measure and the risk premiums ability to explain its movements deserves some further comments. Here, our ambition has been to separate the investment decision by the limited partner and the general partner to capture the scenario where a fund is raised prior to investment considerations by the general partner takes place. As mentioned in Section II.F, our data prevents us from determining the exact timing of fundraising but the statistical significance of the results, for a variety of lags on the risk premiums, highlights the potential importance of the aggregate risk premium in the decision by limited partners to invest in private equity. One should nevertheless keep in mind that our data is far from perfect, especially considering the last couple of years in our sample period. In order for a fund to be included in our regressions, a first investment by the fund has to be made. Funds raised recently, where the first investment is yet to have been announced are thus excluded, making the data incapable of accurately capture the full extent of fund dry powder present in the private equity industry today. The presence of record level dry powder in combination with a low U.S. risk premium does however pose interesting questions regarding the relationship between the equity risk premium and LBO activity in recent years. It could be that the inputs used to forecast future returns applied in this paper does not reflect the risk premium general partners now consider. Alternatively, general partners have not relied on the equity risk premium post-2011 to the same extent as previous years. It would of course be naive to expect that the equity risk premium always explain LBO activity perfectly even if one could estimate it with a high degree of confidence on its accuracy. There are periods earlier in our sample period where the risk premium and LBO activity does not mirror each other perfectly and factors other than the risk premium, such as credit market factors, have been shown to impact the buyout decision. Nonetheless, looking at all evidence put together, our fundraising tests supplement the LBO activity tests and emphasize the role the aggregate risk premium plays explaining the booms and busts of private equity.

VI. Concluding Remarks and Directions for Future Research

By examining a sample of 375 LBOs in the U.K. between 1998Q2 and 2016Q4 and a sample of 1331 LBOs in the U.S. between 1982Q4 and 2016Q4, we aim to further investigate the role the aggregate risk premium play on buyout activity and its relationship to target betas. In the U.S., an extended sample period compared to previous literature captures a time where dry powder levels are higher than ever before, shedding light on the dynamics between the risk premium, buyout activity and target betas post-2011. For the U.K. we test if previous findings remain true in a new setting. The unique patterns for buyout activity in the U.K., with a distinct peak in activity in 1999-2000 provide an interesting testing ground for the relationship between the risk premium and LBO activity and emphasize that the definition of activity itself warrants serious consideration. Furthermore, we investigate new links between how unlevered target betas have developed over time and the link between the risk premium and private equity fundraising. Our results indicate that the relationship between activity and the risk premium is dependent on how and when one estimates the risk premium. This highlights the importance of carefully considering the input variables used when estimating the risk premium. Bringing all evidence together, our results from the U.K. LBO market supports the evidence found in the U.S. A strong link appears to exist between LBO activity and the aggregate risk premium. We do, however, find contrasting results compared to Haddad et al. (2017) regarding the link between the aggregate risk premium and buyout activity on LBO target betas in the U.K. When adding data post-2011 and applying a risk premium with a longer estimation window, there still seems to be a strong link between buyout activity and the risk premium in the U.S. while the risk premium seems to lose its predictive ability on target betas. For both samples, our results indicate that target betas have increased over time.

Regarding the link between fundraising and the aggregate risk premium, our tests indicate that this link might be as strong as investment activity considering only the decision by limited partners to invest in private equity. The literature aiming to explain LBO activity is however far from being exhausted and the results found in this paper leave interesting avenues for future research, a few of which will be discussed in this section.

First, as previously emphasized, an underlying assumption needed to rationalize the relationship between LBO activity and the aggregate risk premium is that general partners and limited partners rely on it to a high extent when deciding to acquire an LBO target. Exactly when and to what extent still remains unknown, however. The survey conducted by Gompers et al. (2016) indicates that private equity firms, at least explicitly, do not rely on valuing projects with Net Present Value rules. Further research into how the risk premium comes into play might therefore be warranted.

Second, perhaps with a better understanding of how private equity firms consider the risk premium, some questions regarding the currently record high levels of dry powder remain unanswered. In a low interest rate and risk premium environment, the low number of LBO deals in recent years is surprising. Investigating the key determinants of this phenomenon would expand our current understanding of the drivers of the booms and busts of private equity. As the private equity industry develops over time, other factors apart from credit measures and the risk premium might prove important.

Third, since we find that our results differ depending on whether PCE or NDS *cay* is used, future research might provide some guidance on which of the two alternative expenditure measures actually capture consumption most accurately. In this paper we are left with the insight that when used to estimate a risk premium aiming to explain LBO activity, the choice of NDS and PCE *cay* influence the results.

The apparent shift in the size of target betas over time also leaves some unanswered questions. We are not any wiser yet regarding why betas have increased although some theories can be formulated. Both the possibility that there has been an overall shift towards higher beta industries and that higher beta firms, regardless of industry, have become more popular LBO targets, remain plausible.

Finally, our fundraising data lacks the precision needed to draw any definite conclusions regarding the relationship between fundraising and the equity risk premium and the relative importance of it in limited partners and general partners investment decisions. More precise data allowing the exact timing of limited partners investments would further help answer when and how the risk premium actually dictates LBO activity.

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Appendix A. Additional Results

Table XV
Drivers of U.K. LBO Weekly Target Betas

This table presents coefficient estimates from regressions where the dependent variable is the unlevered target betas in our U.K. LBO sample based on weekly returns with a two-year estimation window and the independent variables are two activity measures and an U.K. aggregate risk premium. Scaled activity is the number of U.K. LBOs scaled by the number of public firms on the London Stock Exchange the prior quarter. Log transaction value is the log of the sum of transaction value of all LBO deals in our U.K sample each quarter. RP IS, NDS *cay* is the risk premium estimated in-sample with the NDS measure of *cay*, the D/P ratio and three-month U.K. government bond rate. Newey-West standard errors (4 lags) in parentheses.

Dep. Var.: Unlevered Weekly Target Betas

	(1)	(2)	(3)
Scaled activity	-0.0027***		
	(0.0009)		
Log transaction value		0.0197	
		(0.0203)	
RP IS, NDS *cay*			0.0024
			(0.0025)
Number of obs	263	263	263
R-squared	0.027	0.008	0.004

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

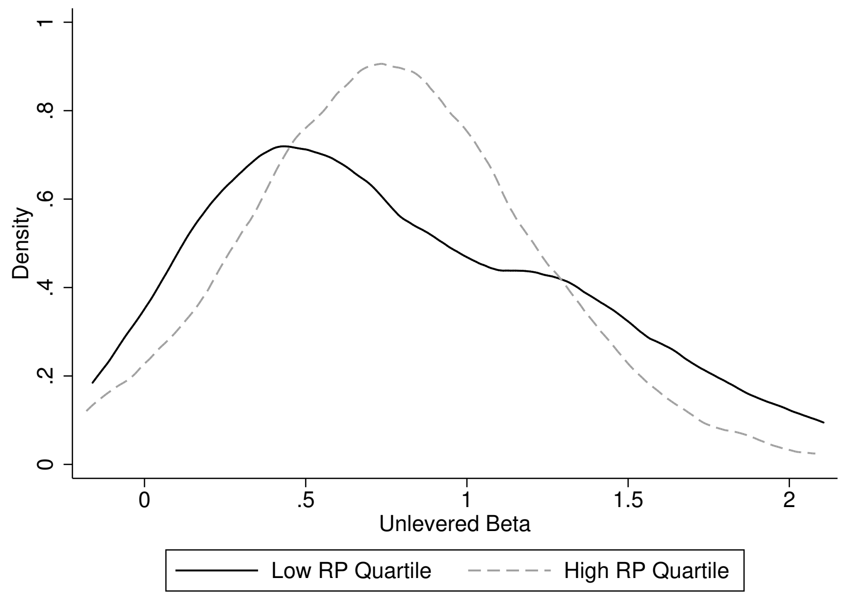

Figure 4. Density of U.S. target betas, high and low IS risk premium quartiles with NDS *cay*. This figure shows how the kernel density of U.S. target betas of the top and bottom risk premium quartiles when the NDS *cay* measure is used in an in-sample risk premium estimation.

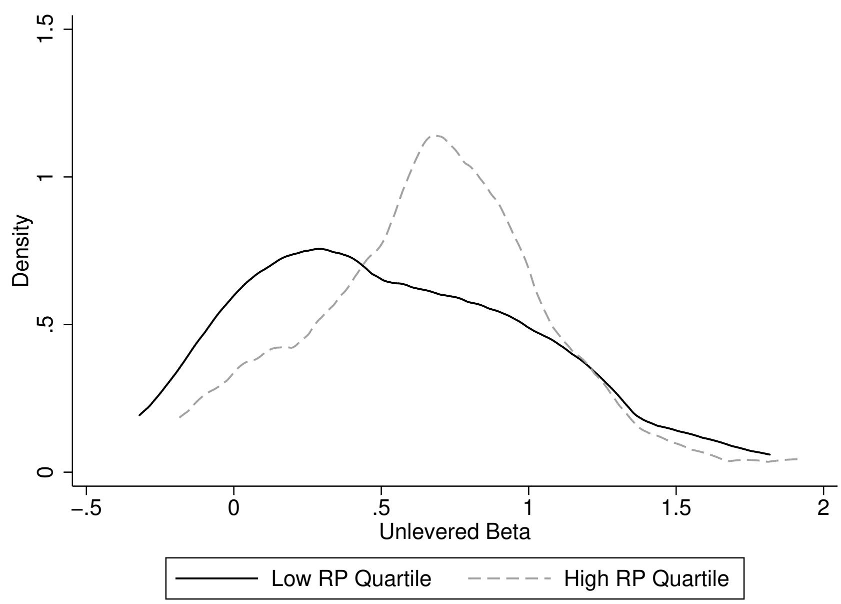

Figure 5. Density of U.K. target betas, high and low IS risk premium quartiles with NDS *cay*. This figure shows how the kernel density of U.K. target betas of the top and bottom risk premium quartiles when the NDS *cay* measure is used in an in-sample risk premium estimation.

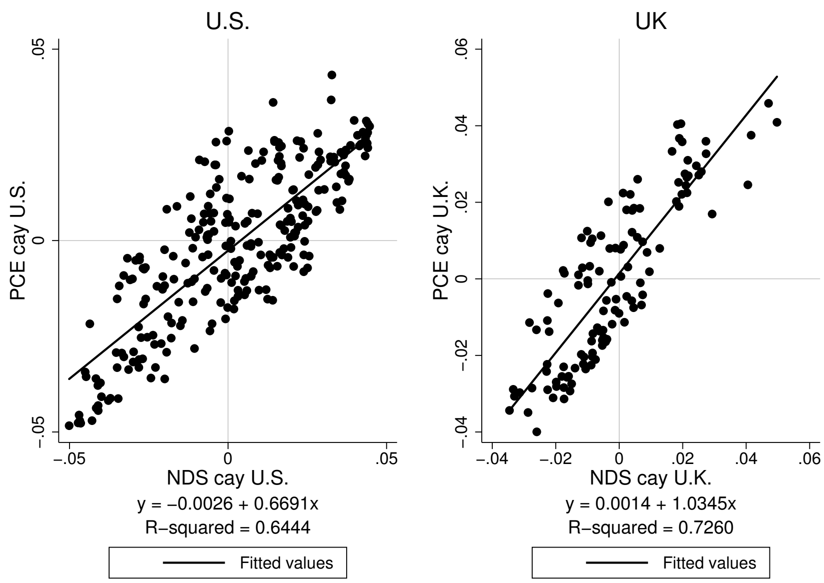

Figure 6. Relationship between NDS *cay* and PCE *cay* for the U.S. and the U.K. This figure shows the relationship between NDS *cay* and PCE *cay* for the U.S. and the U.K. respectively. For the U.S. the estimation period for *cay* is between 1952Q1 and 2016Q3 using dynamic least squares with 8 leads and lags. For the U.K. the estimation period for *cay* is between 1987Q1 and 2016Q3 using dynamic least squares with 1 lead and lag.

Appendix B. In-sample Risk Premium Regression Coefficients

*U.S. NDS cay**

Estimation window 1954Q1-2010Q3.

$$E\left(R^e_{M,t+1}\right) = 0.72 + \frac{(3.34)}{[.48]}(D/P)_t + \frac{(2.54)}{[.28]}cay_t + \frac{(-1.27)}{[.18]}(\mathit{T\text{-}Bill})_t, \tag{B1}$$

U.S. NDS cay

Estimation window 1952Q1-2013Q4.

$$E\left(R^e_{M,t+1}\right) = .41 + \frac{(3.98)}{[.45]}(D/P)_t + \frac{(1.38)}{[.22]}cay_t + \frac{(-1.64)}{[.17]}(\mathit{T\text{-}Bill})_t, \tag{B2}$$

U.S. PCE cay

Estimation window 1952Q1-2013Q4.

$$E\left(R^e_{M,t+1}\right) = 2.47 + \frac{(2.97)}{[.41]}(D/P)_t + \frac{(2.47)}{[.23]}cay_t + \frac{(-1.28)}{[.15]}(\mathit{T\text{-}Bill})_t, \tag{B3}$$

U.K. NDS cay

Estimation window 1987Q1-2014Q1.

$$E\left(R^e_{M,t+1}\right) = -22.19 + \frac{(9.84)}{[.94]}(D/P)_t + \frac{(-.27)}{[.31]}cay_t + \frac{(-1.48)}{[.20]}(\mathit{T\text{-}Bill})_t, \tag{B4}$$

U.K. PCE cay

Estimation window 1987Q1-2014Q1.

$$E\left(R^e_{M,t+1}\right) = -23.01 + \frac{(10.13)}{[.91]}(D/P)_t + \frac{(-.65)}{[.36]}cay_t + \frac{(-1.53)}{[.19]}(\mathit{T\text{-}Bill})_t, \tag{B5}$$

Appendix C. Data Used to Construct the U.K. *cay*

Consumption

Data on expenditures used to proxy for consumption is collected from the Office for National Statistics (ONS), from the report United Kingdom Economic Accounts (UKEA) published on June 30, 2017 for the consumption measure similar to PCE. For the consumption measure similar to NDS, the additional data series are collected from the Office for National Statistics webpage. We define two different measures of consumption. One similar to the U.S. PCE consumption (personal consumption expenditure) and one similar to NDS (nondurable goods and services). The data for both consumption measures is in its original form in current prices, seasonally adjusted and in millions of pounds. We adjust it to 2001Q4 prices expressed in per capita terms and in logarithmic form. Our measure for PCE is defined as Individual consumption expenditure for Households and Non-Profit Institutions Serving Households. For our NDS consumption measure, we follow Sousa (2010) where the consumption of Durable and Semi-durable goods is subtracted from Individual Consumption expenditure for Households and Non-profit Institutions Serving Households.

Labor Income

For labor income, we follow Della Corte et al. (2010) and consider disposable income as a proxy for labor income. Data is again collected from ONS and the UKEA report published on June 30 2017. Labor income is defined as Gross Disposable Income for Households and Non-Profit Institutions Serving Households. The data in its original form is in current prices, seasonally adjusted and in millions of pounds. As we do for consumption, we adjust it to 2001Q4 prices expressed in per capita terms and in logarithmic form.

Financial Wealth

For financial wealth, data is collected from the ONS and yet again from the UKEA report published on June 30 2017. We follow Sousa (2010) and consider Net Financial Assets for Households and Non-Profit Institutions Serving Households. The data in its original form is in current prices, not seasonally adjusted and in millions of pounds.

Housing Wealth

For Housing wealth, we use the Datastream data series on dwellings (UKCGRI..), available only at an annual basis. Since house prices is the prime driver for housing wealth, we interpolate house prices to a quarterly basis using the Halifax housing index suggested by Elbourne (2008) to be the most appropriate housing index to use for the U.K. market. Quarterly interpolation is made through readjusting the annual growth with the proportional quarterly growth in house prices according to the index. The data of dwellings is in millions of pounds in current prices with no seasonal adjustment.

Total Wealth

The sum of our financial wealth and housing wealth measures constitute total wealth. The data is then adjusted to 2001Q4 prices, expressed in per capita and in logarithmic form. The wealth data was tested for seasonality using the X-13-ARIMA software, the industry standard for seasonal adjustment available from the U.S. Census Bureau, with no distinct effect on the numbers.

Population

We again follow Sousa (2010) and population data is gathered from ONS. We define population as the mid-year estimates available (DYAY). To find the quarterly estimates we interpolate between mid-year estimates by computing the annual population growth and then adjusting it to quarterly rates. The series comprise the period 1986Q2-2016Q2. For the last quarter, 2016Q3, the same quarterly growth rate as in between 2015Q2-2016Q2 is applied.

Price Deflator

Nominal prices are deflated by the All-Items Retail Price Index gathered from the Office for National Statistics. The data is quarterly but the time values are adjusted according to annual rates. The index is adjusted so that index=100 occur at 2011Q4.

Comparability with Existing Literature

Sousa (2010) constructs a U.K. *cay* estimate with data from 1980Q1 to 2007Q4 and reports a co-integration estimate for asset wealth of 0.17 and 0.75 for after-tax labor income respectively. Our construction of PCE *cay* estimated from 1987Q1 to 2016Q4 has co-integration estimates of 0.15 for asset wealth and 0.73 for labor income.

Office for National Statistics Series Used in the Estimation of the Consumption Measures

PCE estimate - ZAKV
All-Items Retail Price Index - CHAW (with 2001Q4=100)
Population - UKPOP
Durable Goods - UTIB
Semi-durable goods - UTIR
Disposable income - RPHQ
Financial net worth - NZEA

Appendix D. Data Used to Construct the U.S. *cay*

PCE

Our PCE *cay* is constructed as described in the paper by Lettau and Ludvigson (2001) apart from the definition of consumption used (the 2001 paper uses NDS consumption) and the prices index applied. The current estimates of PCE, which can be found at Martin Lettau's webpage, uses 2009 as base year and we do too. Our measure of PCE is therefore Personal Consumption Expenditures per capita in 2009 chain-weighted dollars. The series for Personal Consumption Expenditure is in its original form seasonally adjusted and in billions of dollars. This is deflated with the PCE chain-price inflator, with 2009=100 which is seasonally adjusted. The measure is expressed in per capita terms through dividing the earlier expression with population. The series for population collected from FRED is expressed in thousands and therefore must be multiplied accordingly to end up with the correct figures. The natural logarithm is finally applied to the numbers each quarter.

NDS

NDS *cay* is constructed as described in the paper by Lettau and Ludvigson (2001) with one exception where we change the base year of the price indices to 2009. NDS is defined as expenditure on nondurable goods and services excluding clothing and footwear. The three series, nondurable goods, services, clothing and footwear are collected separately and are in their original form seasonally adjusted and in billions of dollars. The three series are deflated by their separate price-index where all have 2009=100 and is seasonally adjusted. NDS is then calculated by adding nondurable goods and services and subtracting clothing and footwear. The same per capita treatment as for PCE is applied and the natural logarithm is finally applied to the numbers each quarter.

FRED Series Used in the Estimation of the Consumption Measures

PCE - PCEC
PCE price index - PCECTPI (with 2009=100)
Population - B230RC0Q173SBEA
The code used for expenditure of nondurable goods is PCND
Expenditure on services - PCESV
Expenditure on clothing and footwear - DCLORC1Q027SBEA
Price index on expenditure of nondurable goods - DNDGRG3Q086SBEA (with 2009=100)
Price index on expenditure of services - DSERRG3Q086SBEA (with 2009=100)
Price index on expenditure on clothing and footwear - DCLORG3Q086SBEA (with 2009=100)