Is there Really no Conglomerate Discount?*

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Abstract

Recent research questions the finding of a conglomerate discount. Most prominently, the diversification decision has been shown to be determined endogenously. When this endogeneity is accounted for, there is no evidence of a conglomerate discount. In addition, it is argued that the risk-reducing effect of diversification increases the value of debt and, as a consequence, firm value for diversified firms is underestimated when the book value of debt is used. However, we argue that the potential effect of accounting for differences between the market and book value of debt on the conglomerate discount is very limited and underscore this conjecture empirically. We also investigate the importance of the estimation technique and show that the neglect of firm fixed effects may erroneously lead to the conclusion that there is no conglomerate discount. When we account for fixed effects, the conglomerate discount remains statistically and economically significant – even in an instrumental variables or Heckman selection-model framework. As prior studies are restricted to the pre-1998 period when segment reporting was under SFAS 14, we perform an out-of-sample test by investigating the post-1997 period after the introduction of SFAS 131, which is less prone to segment undereporting.

Keywords: Organizational structure; Diversification; Segment reporting standard; Endogeneity; Fixed effects

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1. Introduction

Recent research questions the finding of a conglomerate discount as reported by Lang and Stulz (1994), Berger and Ofek (1995), and many others. Most prominently, Campa and Kedia (2002) and Villalonga (2004a) show that corporate diversification strategies are determined endogenously and when this endogeneity is accounted for, there is no evidence of a conglomerate discount. In another well-know study, Mansi and Reeb (2002) argue that the risk-reducing effect of diversification results in better credit ratings and higher market values of debt. Consequently, the use of the book value of debt in the calculation of firm value may result in an under-estimation of firm value for diversified firms. In this study, we show that by accounting for firm fixed effects, we find a robust and significant conglomerate discount even when the endogeneity of the diversification decision is accounted for in an instrumental variables framework or Heckman's self-selection model. In addition, we show that the argument of Mansi and Reeb (2002) is implausible and their results are most likely to be caused by a sample selection bias. Most importantly, we find a significant and robust conglomerate discount in the magnitude of approximately 5% to 21% – depending on the regression specification and estimation technique.

The first contribution of this paper is to provide an out-of-sample test on the findings of earlier research by investigating the post-1997 period. The sample period in prior studies on the conglomerate discount usually ends in 1997 or earlier, even for recent studies. For example, the sample periods in Campa and Kedia (2002), Villalonga (2004a), and Ahn, Denis, and Denis (2006) end in 1996, 1997, and 1997, respectively.[1] The reason is that in 1997 the Statement of Financial Accounting Standards (SFAS) 131 superseded SFAS 14 in the regulation of segment reporting (FASB, 1997). At the same time, the Standard Industry Classification (SIC) system was replaced by the North American Industry Classification System (NAICS). One of the major concerns that triggered these changes was an under-reporting of segments. In fact, Berger and Hann (2003) show that

[1] To our best knowledge, there is only one study on conglomerate discounts in non-financial firms using post-1997 data: Mansi and Reeb (2002) use a data sample from 1988 to 1999. In addition, Laeven and Levine (2007) and Schmid and Walter (2007) use post-1997 data in their studies on the valuation effect of corporate diversification strategies in banks and financial intermediaries, respectively.

the implementation of SFAS 131 has resulted in a greater number of segments being reported by at least certain firms. Hence, segment information under SFAS 131 is presumably more accurate and less subject to the concerns of segment under-reporting as raised by Lichtenberg (1991) and Villalonga (2004b). However, as a consequence of this change, segment data before and after 1997 is not directly comparable.

This paper provides an out-of-sample test of the results in prior empirical studies on the conglomerate discount and investigates whether the results hold when segment reporting is based on the more disaggregated SFAS 131. In addition, we investigate how SFAS 131 affects segment reporting and the conglomerate discount. Consistent with Berger and Hann (2003), we find a substantial increase in the percentage of diversified firms in the sample from roughly 18% in 1997 to 27% in 1998 while this percentage number is monotonically decreasing in the years before 1998 and very stable in the years after 1998. In addition, of the 2,795 firms which are in our sample in both 1997 and 1998, 378 firms (14%) increase the number of segments while only 99 firms (4%) report less segments in 1998 than in 1997. Interestingly, firms reporting more segments under SFAS 131 than SFAS 14 experience a substantial drop in excess value from 1997 to 1998 when the "hidden" diversification is revealed. This result holds when we exclude firms undertaking acquisitions in 1998 in order to obtain a "cleaner" reporting change sample. When we estimate the conglomerate discount in a standard pooled OLS regression framework as employed in prior research (e.g., Berger and Ofek, 1995; Campa and Kedia, 2002), the discount is similar in the pre- and post-SFAS 131 period (13% vs. 11%). Hence, although we find evidence for segment under-reporting before the introduction of SFAS 131, the documented conglomerate discount seems to be largely unaffected by these changes in segment reporting.

As a second contribution of this paper, we attempt to investigate whether a potential risk-reducing effect of diversification affects the conglomerate discount. Theory suggests that the combination of business segments with imperfectly correlated earnings streams increases debt capacity through a risk-reducing effect which may result in better credit ratings and higher market values of debt (e.g., Lewellen, 1971). Hence, while diversification may reduce shareholder value, it should enhance bondholder value due to a reduction in firm risk. In fact, two recent studies assess the conglomerate discount in a contingent claims framework and suggest that the use of the book value of debt in the calculation of the excess value measure may lead to a downward bias for diversified firms (Mansi and Reeb, 2002; Ammann and Verhofen, 2006). Mansi and Reeb (2002) underscore this conjecture empirically and show that the use of book value of debt to compute excess values leads to a downward bias for diversified firms. Specifically, when they include the market value of debt instead of the book value of debt in the calculation of the Berger and Ofek (1995) excess value measure, they find no significant diversification discount on average.

However, we argue that organizational structure only affects debt value when the degree of diversification increases or decreases. If a firm's degree of diversification remains unchanged, bonds are issued at par and their value does not depend on diversification. Hence, we argue that differences in debt value can only emerge when a firm changes its degree of diversification. However, such effects on debt value will prevail only for a limited period of time as all subsequent bond issues will again be at par. Consequently, the potential effect of risk-reducing diversification on the conglomerate discount is very limited and we expect it to explain only a small part of the discount if anything.

We investigate this conjecture empirically and use an approach proposed by Damodaran (2005) to estimate the market value of debt. We then include this estimate in the sales- and asset-based excess value measure instead of the book value of debt. Our methodology treats the entire long-term debt on the books as one coupon bond with the coupon set equal to the interest expenses on all debt. We then value this coupon bond at the current cost of debt for the company approximated by the yield of a bond portfolio with the same credit rating. As Compustat provides an official credit rating from S&P only for a very limited subset of our sample, we alternatively construct an artificial credit rating based on the interest coverage ratio. Hence, in contrast to Mansi and Reeb (2002), our approach does not require bond prices for all outstanding bonds of the sample firms.

Our results reveal that by taking into account the market value of debt in the calculation of the excess value measure, the discount is only slightly reduced and remains significant. Hence, the results are consistent with our conjecture that the potential risk-reducing effect of diversification on the market value of debt is very limited. Only when we alternatively construct the market value of debt based on the official credit rating provided by S&P, the discount turns into a significant premium in all regression specifications. However, we show that this result is due to the sample selection process as the credit rating from S&P is only available for 26% of the sample firms.

We also investigate the effect of leverage on the conglomerate discount. Mansi and Reeb (2002) argue that leverage is important as the risk-reducing effect of diversification is expected to increase debt value (while reducing shareholder value) and hence the expected downward bias resulting from using the book value of debt in the calculation of the excess value measure should increase in leverage. While our previous results reject Mansi and Reeb's (2002) wealth transfer story, we also find the conglomerate discount to increase in leverage while there is no conglomerate discount for all-equity firms. However, this result is consistent with alternative explanations as well. For example, high-leverage firms might be in a poor shape and diversify into other activity areas to improve growth opportunities and/or decrease their credit risk with the objective of reducing the costs of debt.

The third and most important contribution of our paper is related to methodological issues in estimating the conglomerate discount. Based on a robust version of the Hausman (1978) test, we substantiate the presence of firm fixed effects. When we account for firm fixed effects, we find a significant conglomerate discount in all regression specifications including those interacting the diversification dummy variable with leverage as explained in the last paragraph. Prior literature suggests that corporate diversification strategies are determined endogenously (Campa and Kedia, 2002; Villalonga, 2004a) and that when this endogeneity is accounted for, the conglomerate discount disappears. When we control for the endogeneity of the diversification decision by estimating similar instrumental variables regressions and Heckman's (1979) self-selection model as used by Campa and Kedia (2002), the results confirm those of Campa and Kedia (2002) in that there is no diversification discount. However, when include firm fixed effects in the second step of the instrumental variables regressions and Heckman's self-selection model, the results confirm the finding of a significant conglomerate discount in the range of 11% to 21%.

The remainder of the paper is organized as follows. Section 2 describes the data, sample selection procedure, and main variables. Section 3 investigates the consequences of the introduction of SFAS 131 on segment reporting and the conglomerate discount. Section 4 contains the analysis of potential risk-reducing effects of diversification and their effect on the calculated excess value measure. Section 5 investigates the importance of the estimation technique and addresses potential endogeneity concerns associated with the diversification decision. Section 6 considers the excess value impact of changes in diversification. Finally, Section 7 concludes.

2. Sample selection and variables

2.1 Sample selection

The sample consists of all firms with data reported on both the Compustat Industrial Annual and Segment data files and covers the period from 1998 to 2005. Following Berger and Ofek (1995) and others, we exclude firm-years in which at least one segment is classified as being in the financial sector (SIC 6000-6999; NAICS 520000-529999) and firm-years with total sales of less than $20 million. Additionally, we exclude firms that are listed as American Depository Receipts (ADRs).

To examine whether diversification increases or decreases corporate value, we rely on the excess value measure developed by Berger and Ofek (1995). For a firm to be included in our sample, all data necessary to calculate this excess value measure are required (see description below). In order to compare our results to those obtained in prior studies using pre-1998 data, we additionally collect data for the period from 1985 to 1997 and apply the same sample selection criteria.

2.2 Measures of excess value

To examine whether diversification increases or decreases corporate value, we use the excess value measure developed by Berger and Ofek (1995) that compares a firm’s value to its imputed value if its segments were operated as stand-alone entities. First, we calculate the imputed value for each segment by multiplying the segment's sales (assets) by the median ratio of market value to sales (assets) for single-segment firms in the same industry. The industry median ratios are based on the narrowest NAICS grouping that includes at least five single-segment firms with complete data and total sales of at least \$20 million.[2] Next, the imputed value of the firm is calculated as the sum of the imputed segment values. Finally, the actual excess value measure is calculated as the log of the ratio of a firm's value to its imputed value. A negative excess value indicates that a firm trades at a discount and a positive excess value implies that the firm trades at a premium.

Some of the segments of diversified firms in our sample have no NAICS codes assigned by Compustat. In contrast, most have a segment name, usually stated as "corporate and other", "eliminations", "corporate and unallocated", or a similar designation. We do not treat these segments separately, but rather attribute their sales (assets) proportionally to the remaining segments in order to sum to the correct figure for the firm's total sales (assets). Nevertheless, for some of the diversified firms in our sample the sum of all segment sales (assets) as provided by the Compustat Segment file disagrees with the respective firm total values from the Compustat Industrial Annual file. We exclude all observations for which the sum of the segment values deviates from the firm's total value by more than 5%. If the absolute deviation is less than 5%, we scale the firm's imputed value up or down by the percentage deviation between the sum of its segments' sales (assets) and total firm sales (assets). Finally, again following Berger and Ofek (1995), we exclude extreme excess values from the analysis. Specifically, we drop all observations where the actual firm value is either larger than four times the imputed value or less than one fourth of the imputed value.

2.3 Measures of diversification

We use a series of alternative measures of diversification. The first is a dummy variable that is equal to one if a firm reports more than one segment in Compustat's Segments

[2] Using sales multipliers, the imputed value for 68.1% of all segments are based on six-digit NAICS codes, 9.0% on five-digit NAICS codes, 12.5% on four-digit NAICS codes, 9.4% on three-digit NAICS, and 1.0% on two-digit NAICS codes. The figures for asset multipliers are basically identical.

data file. Earlier evidence (e.g., Lang and Stulz, 1994) suggests that firms with two or more segments have a lower firm value than firms with one segment, but that there is no further significant drop in firm value when the number of segments increases from j to $j + 1$ segments, where $j \geq 2$. To investigate whether this finding is also valid for our more recent sample, we alternatively use the number of segments reported by Compustat. Additionally, we follow Lang and Stulz, (1994), Comment and Jarrell (1995), and Denis, Denis and Sarin (1997) in using a sales- and asset-based Herfindahl-Hirshman index (HHI). These HHIs are computed as the sum of the squares of each segment's sales (assets) as a proportion of the square of total sales (assets) for the firm. For example, if a firm has only one segment, its HHI is equal to one. If a firm has 10 segments that each contribute 10 percent of the sales (assets), its HHI is equal to 0.1. Hence, the HHI decreases as the degree of diversification increases.

Berger and Ofek (1995) show that only unrelated diversification (i.e., diversification at the two-digit SIC code level) is associated with a significant conglomerate discount while there is no discount for related diversification (i.e., diversification at the four-digit SIC level). We therefore investigate potential differences in the valuation effects associated with related and unrelated diversification. Specifically, we construct a dummy variable which is equal to one if a firm reports more than one segment based on three-digit-level NAICS codes to measure unrelated diversification, and a similar dummy variable, which has a value of one if a firm reports more than one segment based on five-digit-level NAICS codes to measure related diversification.

3. Is there a Conglomerate Discount in the Post-1997 Period?

3.1 Descriptive analysis

This section provides an out-of-sample test of the results in prior empirical studies on the conglomerate discount and investigates whether the results hold when segment reporting is based on the more disaggregated SFAS 131. First, we conduct a univariate analysis and investigate the evolution of the percentage of diversified and focused firms as well as the discount associated with running a multi-segment business over time. Table 1 reports the number of total observations, the percentage of focused firms in the sample, the percentage of diversified firms in the sample, and the mean value of the sales-based excess value measure for focused and diversified firms and each sample calendar year. All firms reporting more than one segment (with differing SIC or NAICS) codes are classified as diversified.

Consistent with prior studies based on the pre-1998 sample (e.g., Comment and Jarrell, 1995; Denis, Denis, and Yost, 2002), our results indicate a steady trend toward greater focus over the period from 1985 to 1997. Specifically, the percentage of diversified firms drops monotonically from 39.0% in 1985 to 18.0% in 1997. By contrast and consistent with Berger and Hann (2003), the reporting change in 1998 leads to a substantial increase in the percentage of diversified firms in the sample from 18.0% in 1997 to 26.5% in 1998. Hence, segment information under SFAS 131 is in fact more disaggregated and mitigates concerns about segment under-reporting.

Looking at firm valuation, Table 1 shows that the diversification discount is statistically significant at the 1% level during the complete pre-1998 period and ranges from 6.6% to 18.0%. Between 1998 and 2002 the discount is substantially reduced (with the exception of 2001) and then becomes large and significant once again in the last three sample years from 2003 to 2005. Alternatively, we investigate the excess value measure based on total assets for each sample calendar year. The results remain qualitatively similar and are not reported in a table for brevity. Most importantly, the discount becomes somewhat higher in the post-1997 period and is statistically significant in all years at the 10% level or better. In addition, we calculate the median for both excess value measures and each sample year, whereas the equality of medians is tested based on a Wilcoxon signed rank test. The respective results remain basically unchanged as compared to those for the means (not reported in a table) suggesting that the results in Table 1 are unlikely to be caused by outliers.

In Table 2, we check the robustness of the results in Table 1 by estimating cross-sectional regressions of the excess value measure on a dummy variable whether the firm is diversified and the standard set of control variables as proposed by Berger and Ofek

(1995) for each sample calendar year. Panel A reports the results for the excess value measure based on sales and Panel B for the excess value measure based on assets. The results in Panel A reveal that when firm characteristics such as size (measured by the log of total assets), capital expenditures (standardized by sales), and profitability (EBIT to sales) are accounted for in a multivariate framework, we find a significant and substantial conglomerate discount for each sample year in the post-1997 period (with the lowest value reported for 2000). The results for the excess value measure based on assets in Panel B are qualitatively similar but even stronger. Here the conglomerate discount ranges from 7.9% (2000) to 16.7% (2005) and is statistically significant at the 1% level in each sample year.

3.2 What happens between 1997 and 1998 upon the introduction of SFAS 131?

In this section, we investigate the changes in the number of segments, excess value, and the sample composition occurring between 1997 and 1998, when SFAS 14 was replaced by SFAS 131. Panel A of Table 3 reports the number of observations and the sales-based excess value measure for 1) previously focused firms which become diversified in 1998, 2) diversified firms increasing the number of segments from 1997 to 1998, 3) diversified firms decreasing the number of segments between 1997 and 1998, and 4) previously diversified firms becoming focused in 1998. Most importantly, 378 firms report more segments based on SFAS 131 as compared to SFAS 14 (14% of all 2,795 firms which are in the sample in 1997 and 1998) while only 99 (4% of the firms) report less segments. This finding is consistent with Berger and Hann (2003) and confirms that SFAS 131 induces segment reporting on a more disaggregated and presumably more appropriate level. Interestingly, the results in Panel A also show that increases in the number of segments are associated with a substantial drop in excess value while decreases in the number of segments are accompanied by a reduction of the discount. Unfortunately, it is difficult to separate changes in the number of segments that are caused by “real” diversification activities and changes caused by the reporting change. As 72 firms undertake acquisitions amounting to 10% of their sales or more in 1998, we exclude these observations to obtain a “cleaner” reporting change sample. The results show that the 254 firms becoming diversified with no or only a relatively small acquisition experience a drop in excess value which is comparable to the numbers reported in Panel A for all firms: from a mean (median) of -0.07 (-0.01) to -0.19 (-0.22). In contrast, firms undertaking acquisitions experience no decrease in firm value at all and still report a premium in the magnitude of 12% (13%) in 1998. Similarly, the 40 diversified firms increasing the number of segments without an acquisition experience a substantial drop in excess value while the 12 firms untertaking at least one acquisition in 1998 show only a slight decrease in the premium. Hence, our results indicate that firms reporting more segments under SFAS 131 and thereby revealing their "true" (or at least more appropriate) level of diversification, experience a substantial increase in the discount. These results are consistent with those of Berger and Hann (2003), who investigate the change in the diversification discount in a hand-collected sample of 543 firms reporting more than one segment in 1998 based on SFAS 131 as compared to only one segment in 1997 under SFAS 14. Specifically, they find a slight discount already in 1997 indicating that the market is partially able to see through the data reported under SFAS 14 but a substantial increase in the discount when the "hidden" diversification is revealed in 1998 under SFAS 131.

Panel B shows the mean and median values of the excess value measure based on sales for the 1,977 focused firms and the 341 diversified firms which are in the sample in 1997 and 1998 and experience no change in the number of reported segments. The mean and median excess value of the diversified firms remains basically unchanged while that of the focused firms slightly decreases which is probably due to the 145 relatively high-valued focused firms entering the sample in 1998 as reported in Panel C.[3,4]

3.3 Multivariate pooled regressions

In this section, we test for the existence of a conglomerate discount in the post-1997 sample and estimate pooled OLS regressions comparable to those reported in prior research covering the pre-1998 period. The choice of control variables is based on prior

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[3] Note that the sum of all 1998-firms in Panel A, B, and C of Table 3 add up to 2,966 firms only as compared to 3,285 observations reported in Table 1 for 1998. The remaining 319 observations are due to firms with gaps in 1997 either because of missing observations or a violation of one or more of the restrictions applied in the calculation of the excess value measure.

[4] We alternatively repeat the analysis in Table 3 based on the asset-based excess value measure and find the results to be qualitatively similar. Therefore, we do not report them in a table.

research as well (e.g., Berger and Ofek, 1995) and includes the natural logarithm of total assets (ln(Assets)), the ratio of capital expenditures to sales (CAPX/Sales), and the ratio of EBIT to sales (EBIT/Sales). In addition, we include the past growth in sales over the last three years (Past Sales Growth) to control for growth opportunities (e.g., see Yermack, 1996). Throughout the whole section, we estimate pooled cross-sectional time series regressions with year fixed effects and Driscoll and Kraay (1998) standard errors, which are heteroskedasticity-consistent and robust to general forms of cross-sectional and temporal dependence.[5]

The results in Columns 1 and 2 of Table 4 compare the results for the pre-1998 and post-1997 periods. The results for the pre-1998 period in Column 1 reveal a conglomerate discount in the magnitude of 13.2%, which is very similar to the 13% obtained by Campa and Kedia (2002) for the 1978 to 1996 period or the 14.4% obtained by Berger and Ofek (1995) for the 1986 to 1991 period (both based on the sales-based excess value measure). Most importantly, the results in Column 2 show that the discount is only slightly reduced to 11.5% in the post-1997 period. Not surprisingly, over the full sample period from 1985 to 2005, the discount amounts to nearly 13% (Column 3). In Column 4, we check the robustness of this result to the inclusion of R&D expenditures scaled by sales for the post-1997 period. While the coefficient on R&D expenditures is positive and significant, the conglomerate discount remains basically unchanged. We reestimate the regressions also for the pre-1997 and the full sample period by including R&D expenditures and find the results to be robust.[6] In Columns 5 and 6, we replace the diversification dummy variable by two alternative measures of diversification, the number of segments and a sales-based Herfindahl-Hirshman index. The results are consistent with those in

[5] Driscoll and Kraay (1998) show that erroneously ignoring cross-sectional dependence in the estimation of linear panel models can lead to severely biased statistical inference. Moreover, Hoechle and Zimmermann (2007) show that the calendar time portfolio approach frequently employed in long-term event studies replicates the Driscoll and Kraay covariance matrix estimator for pooled OLS regressions by aid of a two-step procedure.

[6] As the inclusion of R&D expenditures substantially reduces the number of observations we henceforth report results from regressions without R&D expenditures only. However, unreported tests show that the results remain qualitatively similar when R&D expenditures are included. Alternatively, we replace missing values of R&D expenditures by zero and add a dummy variable which is equal to one if R&D expenditures are missing and zero otherwise. Again the results remain qualitatively similar whereas the coefficient on R&D expenditures is in general positive and significant and the coefficient on the dummy variable whether R&D expenditures are missing is negative and significant.

Column 2 and reveal a negative and significant relationship between firm value and the number of segments as well as firm value and the concentration of activities over the different segments. Finally, in Columns 7 and 8, we investigate whether the conglomerate discount is largely restricted to unrelated rather than related diversification in the post-1997 sub-sample as suggested by prior research (e.g., Berger and Ofek, 1995). In fact, the results show that only the coefficient on the dummy variable measuring unrelated diversification is statistically significant. Nevertheless, we use the standard definition of the dummy variable in the remainder of the paper to keep our results comparable to those of Campa and Kedia (2002) and others who also use this definition. However, unreported tests show that our results remain qualitatively unchanged when we use the dummy variable based on unrelated diversification.

4. The Conglomerate Discount and Risk Reduction

As already noted in the introduction, Berger and Ofek (1995) and the vast majority of subsequent studies adopting their methodology rely on using the market value of equity plus the book value of debt as a proxy for a firm's market value. However, Mansi and Reeb (2002) argue that the risk-reducing effect of corporate diversification increases debt value and therefore using the book value of debt to compute the excess value measure creates a downward bias for diversified firms. Mansi and Reeb (2002) underscore their conjecture empirically and show that the diversification discount disappears when they include the market value of debt rather than the book value of debt in the computation of Berger and Ofek's (1995) excess value measure. To calculate the market value of debt, Mansi and Reeb (2002) rely on the Lehman Brothers Fixed Income database. This causes two potential problems: First, data is only available for a very limited subset of the original sample (2,487 out of 18,898 firm-year observations). Second, many firms have non-traded debt, such as bank debt, which is specified in book value terms but not in market value terms. In addition, the database is not maintained anymore and therefore cannot be used for our more recent sample.

The argument of Mansi and Reeb (2002), however, applies only to changes in diversification. If a firm's degree of diversification remains unchanged, bonds are issued at par and their value does not change unless market rates change which, however, is not a function of conglomeration. Hence, differences in debt value can only occur when a firm changes the degree of diversification and will prevail only for a limited period of time as all subsequent bonds issues will be again at par. Consequently, we expect the potential effect on the conglomerate discount to be of minor importance. In addition, Baecker and Grass (2007) calculate the expected conglomerate discount resulting from the risk-reducing effect of diversification in a contingent claims framework. The obtained discount attributable to diversification amounts to 0.9% only for the mean multi-segment firm in their Compustat-sample.[7]

We investigate the potential effect of differences between the market and book value of debt on the conglomerate discount based on an alternative methodology proposed by Damodaran (2005). This approach does not require data on bond prices for all outstanding bonds of the firms. We treat the entire long-term debt on the books as one coupon bond with the coupon set equal to the interest expenses on all debt and the maturity set equal to the face value weighted average maturity of the firm's debt. We then value this coupon bond at the current cost of debt for the company. Thus, the market value of the long-term debt is estimated as follows:

$$Market\,Value\,of\;\,Long-Term\;Debt = Interest\;Expenses\left(\frac{1-\dfrac{1}{\left(1+Current\;Cost\;of\;\,Debt\right)^{Maturity}}}{1-\dfrac{1}{\left(1+Current\;Cost\;of\;\,Debt\right)}}\right)$$

$$+\frac{Book\,Value\,of\;\,Long-Term\;Debt}{\left(1+Current\;Cost\;of\;\,Debt\right)^{Maturity}}\,.$$

To estimate the current cost of debt, we use the yield to a bond portfolio with the same credit rating. As Compustat provides a credit rating (from S&P) for approximately 22% of our sample firms only, we calculate "artificial" bond ratings for each sample firm (and year) based on a firm's interest coverage ratio defined as earnings before interest and taxes divided by interest expenses. The interest coverage ratio measures the number of

[7] As such wealth transfers between share- and bondholders are only temporary (until debt contracts are re-negotiated), this estimate has to be interpreted as an upper limit.

times a company could make its interest payments with its earnings before interest and taxes and therefore proxies for the company's debt burden. To obtain a credit rating for each sample firm, we fit the empirical distribution of the 15,433 firm-year observations (non-financial firms with total sales in excess of $20 million) with an official S&P credit rating on Compustat over our sample period to the calculated interest coverage ratios of the 53,874 firm-year observations with no S&P credit rating. The yields for the corresponding bond portfolios stem from Bloomberg.[8] We set the maturity for this long-term debt equal to 10 years.[9] Finally, we add the book value of short-term debt due in one year to obtain the market value of the firms' debt.

This approach points out that an increase in the credit rating (and the corresponding decrease in the current cost of debt) is associated with a decrease in interest expenses. Consequently, there is a counteracting effect for the valuation of the coupons (interest expenses) which limits the deviations of the market value of debt from the book value of debt. Hence, only unexpected changes in the current cost of debt – such as an improved debt rating due to a diversifying activity which reduces firm risk – affect the value of debt. As a consequence, in contrast to Mansi and Reeb (2002), we expect only a modest reduction in the conglomerate discount when the market value instead of the book value of debt is included in the excess value measure.

Column 1 of Table 5 reports the results from reestimating the regression equation in Column 2 of Table 4 for the excess value measure including the estimate of the market value of debt based on the artificial credit rating. As expected, the results show that the conglomerate discount is slightly reduced but remains significant at the 1% level. In Column 2, we additionally include leverage and an interaction term between leverage and the diversification dummy variable as additional control variables. Mansi and Reeb (2002) argue that as the risk-reducing effect of diversification is expected to increase debt value (while reducing shareholder value), the expected downward bias resulting from using the book value of debt in the calculation of the excess value measure should increase in lev-

[8] We also calculate an excess value measure based on the official credit rating provided by S&P and available on Compustat for the limited sub-sample.

[9] Alternatively, we use six years. However, the results remain qualitatively similar and, therefore, we only report results for a maturity of 10 years in the paper.

erage. In addition, leverage might affect firm value based on the role of debt in helping to discourage the overinvestment of free cash flow by self-serving managers (e.g., Jensen, 1986; Stulz, 1990; Hart and Moore, 1995). Debt can also create value by giving the management an opportunity to signal its willingness to distribute cash flows and to be monitored by lenders. Empirically, McConnell and Servaes (1995) find that book leverage is positively correlated with firm value when investment opportunities are scarce, which is consistent with the hypothesis that debt alleviates the overinvestment problem. Most importantly, the results reveal that the conglomerate discount disappears and turns into a slight but insignificant premium. In addition, the coefficient on leverage is estimated negative and significant while the coefficient on the interaction term between leverage and diversification is negative but insignificant.

In Column 3, we check the robustness of these results with respect to the augmented regression specification as proposed by Campa and Kedia (2002) and include their full set of control variables including the log of total assets, the ratio of capital expenditures to sales, and EBIT to sales (each with one and two lags), and the log of total assets squared. The results show that many of the additional control variables are statistically significant while the diversification dummy variable remains basically unchanged as compared to Column 2. However, the coefficient on leverage becomes insignificant indicating that these additional control variables capture the effect of leverage on firm value.

In Column 4, we include the market value of debt based on the official credit rating provided by S&P in the excess value measure. The results show that the diversification dummy variable turns positive and significant indicating a premium associated with corporate diversification strategies. Moreover, the coefficient on the interaction term between leverage and the diversification dummy variable is estimated negative and significant. This latter finding is consistent with Mansi and Reeb (2002) and indicates that the discount strongly increases in leverage. However, this finding is also consistent with other explanations than the one based on a wealth transfer between debt and equity holders provided by Mansi and Reeb (2002). We propose an alternative explanation at the end of this section. However, by requiring the availability of the S&P credit rating and thereby reducing the sample size by roughly 74% from 16,001 to 4,227 firm-year observations, we might introduce a sample selection bias. In fact, when we reestimate the regression for the sample for which S&P ratings are available but use the excess value measure including the book value of debt, the results show a significant premium as well (see Column 5). Consistently, when we alternatively use the excess value measure including the estimate of the market value of debt based on the artificial credit rating, the diversification dummy variable is estimated to be positive and significant with a coefficient equal to 0.14 and a *t*-value equal to 2.25 (results not reported). Therefore, we conclude that the disappearance of a significant conglomerate discount (and emergence of a significant premium) is likely to be caused by the sample selection process resulting from the requirement of available credit ratings from S&P. Mansi and Reeb (2002) might encounter the very same problem. In fact, their sub-sample including the market value of debt only accounts for roughly 13% of the total sample.[10]

The importance of the sample selection process is confirmed by unreported univariate tests: The sample with S&P ratings available exhibits an average premium for diversified firms (even when the book value of debt is included in the excess value measure) while the mean excess value is between -0.0908 and -0.1157 (depending on the excess value measure) for the larger sample including 17,075 firms with an artificial credit rating. In addition, the sample firms with S&P ratings available differ substantially with respect to various other variables, such as for example firm size (significantly larger) or leverage (significantly higher). Consequently, we henceforth do not report results on the excess value measure including the estimate of the market value of debt based on the official S&P credit rating.

Finally, in Columns 6 and 7, we estimate the role of leverage on the conglomerate discount by estimating the standard regression specification including the excess value

[10] In addition, Worldscope, the database used by Mansi and Reeb (2002), may introduce a survivorship bias as there was no research tape including delisted/dead firms for this database at this time. In fact, the conglomerate discount for the full sample based on the book value of debt (7.1% and 4.5% in univariate and multivariate tests, respectively) is substantially lower than in other studies covering a similar time period. As a comparison, we find a conglomerate discount of 10.6% (10.4%) in univariate (multivariate) tests based on the sales-based excess value measure (including the book value of debt) over the time period from 1990 to 1999 which covers the largest part of Mansi and Reeb's sample period.

measure based on the book value of debt as well as the extended version including the additional control variables as suggested by Campa and Kedia (2002). Consistent with the results in Columns 2 and 3, the diversification dummy variable becomes insignificant. Moreover, consistent with the results in Columns 4 and 5, the coefficient on the interaction term between leverage and diversification is estimated negative and significant. Hence, the discount strongly increases in leverage and there is no conglomerate discount at all for all-equity firms. While this finding is consistent with the results from Mansi and Reeb (2002), we believe that the effect of leverage is due to other reasons than a risk transfer from equity to debt holders in conglomerates. For example, the discount may be more pronounced for high-leverage firms because these firms are in a worse shape – which in turn might be the reason why they are in a conglomerate. Hence, diversification might be endogenous as suggested by Campa and Kedia (2002) and Villalonga (2004a). We account for the endogeneity of the diversification variable in Section 5 by aid of an instrumental variables approach and Heckman's self-selection model.

5. Methodological issues in estimating the conglomerate discount

5.1 Fixed effects regressions

In this section, we start by estimating fixed effects regressions to control for potential omitted variables which are either constant over time or constant over firms. Using fixed effects regressions is also potentially important as we argue that organizational structure only affects debt value when the degree of diversification increases or decreases. Hence, we reestimate the analyses in Table 5 by inlcuding firm fixed effects. The results are reported in Table 6. Most importantly, the results show that the diversification dummy variable is negative and significant in all regression specifications with the exception of Columns 4 and 5 when the sample is restricted to firms with a credit rating from S&P available. For this sub-sample, we again report a significant premium associated with corporate diversification strategies. In Columns 6 and 7, when the excess value measure includes the book value of debt, the conglomerate discount amounts to roughly 12% and is therefore similar to the results reported in Table 4. In addition, the interaction term between leverage and the diversification dummy variable turns insignificant while the coefficient on leverage is positive and significant in Column 7.

We use a Hausman (1978) specification test to test for the presence of firm fixed effects. Although pooled OLS regression yields consistent coefficient estimates when the random effects model is true (i.e., the unobserved effects are uncorrelated with the regressors), its coefficient estimates are inefficient under the null hypothesis of the Hausman test. Because feasible generalized least squares (FGLS) estimation is both consistent and efficient under the null hypothesis of the Hausman test, the coefficient estimates obtained from FGLS should be compared with those of the fixed effects estimator. As the random effects estimator is not fully efficient under the null hypothesis when the unobserved effects or the error term are not i.i.d., we perform the alternative version of the Hausman test based on a Wald test in an auxiliary OLS regression.[11] Wooldridge (2002) and Hoechle (2007) recommend to use this version of the test and to estimate the auxiliary regression with robust standard errors to ensure valid inference also when the unobserved effects or the error term are not i.i.d. To assure that the test is robust to heteroskedasticity and general forms of spatial and temporal dependence, we fit the auxiliary regression with Dricoll and Kraay (1998) standard errors. The results of this robust version of the Hausman test are reported in Table 6 and reveal that the random effects assumption is rejected in all seven specifications and consequently firm fixed effects should be included in all regressions. When we check the robustness of Table 4 with respect to the inclusion of firm fixed effects, however, the results remain qualitatively unchanged.

5.2 The endogeneity of the diversification decision

Recent research suggests that corporate diversification strategies are determined endogenously (e.g., Campa and Kedia, 2002; Villalonga, 2004a). In this case it is not possible to assess causation based on the results of OLS regressions as estimated in Tables 4 and 5 (and potentially also Table 6 if the unobserved heterogeneity that leads to the correlation between the diversification dummy variable and the error term is not constant

[11] Hausman (1978) shows that this alternative specification of the test is asymptotically equivalent to the usual chi-squared test. While this alternative formulation of the test does not necessarily have better finite-sample properties than those of the standard Hausman test, it has the advantage of being computationally more stable in finite samples because it never encounters problems with non-positive definite matrices.

over time or firms). In this section, we first follow Campa and Kedia (2002) and Villalonga (2004a) and account for a potential endogeneity of the corporate diversification strategy by estimating instrumental variables regressions as well as Heckman's (1979) self-selection model. Subsequently, we account for the endogeneity of the diversification decision and the relevance of fixed effects by including year *and* firm fixed effects in the second step of the instrumental variables regressions and Heckman's self-selection model.

We use a Hausman specification test in order to test for the presence of endogeneity (Hausman, 1978). The test is based on a comparison of the estimator from an instrumental variables regression (which is consistent under both the null and the alternative hypotheses but inefficient under the null hypothesis) and the OLS estimator (which is consistent and efficient under the null hypothesis of no endogeneity but inconsistent under the alternative hypothesis). A key issue is the choice of instruments in the instrumental variables regression as many of the natural instruments for the diversification dummy variable are already included in the excess firm value equation. We rely on Campa and Kedia (2002) and Villalonga (2004a) and use a set of firm characteristics and a set of industry and time characteristics. The first set includes a dummy variable whether the firm is listed on NYSE, Nasdaq, or AMEX, a dummy variable whether the firm belongs to the S&P industrial index, a dummy variable whether the firm is incorporated outside the U.S., the log of total assets (with zero, one, and two lags), EBIT/sales (with zero, one, and two lags), CAPEX/sales (with zero, one, and two lags), and the historical average values of the log of total assets, EBIT/sales, and CAPEX/sales. The set of industry and time characteristics includes the fraction of all firms in an industry which are conglomerates and the fraction of sales by other firms in the industry accounted for by diversified firms, the number and value of merger and acquisition announcements in a given year, and real growth rates of the GDP and its lagged value.[12] We use four-digit NAICS codes to identify industries (however, the results are robust to alternative definitions such as three- or five-digit NAICS codes). Independent of the choice of control variables and instruments, we can reject the null hypothesis of no endogeneity at the 1% level which

[12] Data on the number and value of M&A transactions are from Thomson Financial's SDC (Securities Data Corporation) database, and data on GDP growth are from NBER.

leads us to the conclusion that the diversification dummy variable is in fact endogenously determined.

To account for the endogeneity of the diversification variable, we first estimate an instrumental variables regression where the endogenous diversification dummy variable is instrumented. In the first stage, we regress the diversification dummy variable on all presumably exogenous variables in the excess value regression along with the predicted probability of being diversified, which is obtained from a probit regression of the diversification dummy variable on various instruments.[13] We use the same set of instruments in this probit regression as in the Hausman specification test. In the second stage, we regress the excess value measure on the fitted value from the first stage, a number of control variables, and a set of year dummy variables (which are not reported in the table). The results are reported in Column 1 of Table 7. Most importantly, we find the discount to decrease substantially and become insignificant. Campa and Kedia (2002) even report a significant premium for the 1978 to 1996 period when the endogeneity of the diversification dummy variable is accounted for in the same instrumental variables setting. However, Campa and Kedia (2002) investigate the decision to diversify and the decision to focus separately and hence restrict the sample either to single-segment and diversifying firms or to single-segment and refocusing firms. When we introduce this same sample partition, the discount is insignificant for both sub-samples (not reported in a table).

In Column 2, we additionally include firm fixed effects in the second step of the instrumental variables regression as reported in Column 1. Most importantly, the results show that the conglomerate discount increases and turns statistically significant at the 10% level. At the same time, the coefficient on leverage turns positive and significant. In Column 3, we do not impose the (nonlinear) functional form of the probit model and use the standard two-stage least squares approach. Hence, in the first step regression, we di-

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[13] We also estimate two alternative models. The first is based on Campa and Kedia (2002) and uses the predicted probability of diversifying (i.e., a dummy variable which is equal to one when a firm increases the number of segments and zero otherwise) instead of the predicted probability of being diversified in the first-step regressions. In the second alternative model we directly include all exogenous variables and instruments in the first step regressions (instead of using the predicted probability of being diversified (or to diversify)). This latter model does not impose the (nonlinear) functional form of the probit model. However, the results from all different specifications are qualitatively similar and therefore we do not report the results from the alternative specifications in a table.

rectly regress the diversification dummy variable on all exogenous variables and instruments (instead of using the predicted probability of being diversified as in Columns 1 and 2). As compared to Column 2, the conglomerate discount in Column 3 further increases to roughly 20% and is now statistically significant at the 1% level.[14]

Column 4 reports the results from estimating Heckman's (1979) self-selection model. In the first step, we estimate a probit regression with a dummy variable whether the firm is increasing the number of segments as the dependent variable. The choice of explanatory variables is the same as in the first-step probit regression of the instrumental variables approach and the Hausman test and is based on Campa and Kedia (2002). In the second stage, we regress the sales-based excess value measure on the dummy variable whether the firm is diversified, the full set of control variables, and the self-selection parameter (lambda). This lambda parameter (or inverse mills ratio) accounts for the correlation between the error terms in the value equation and the (probit) equation modeling the firms probability to diversify. The results reveal that the coefficient on the diversification dummy variable is again positive and insignificant while the coefficient on the self-selection parameter is negative and significant. This latter finding indicates the prevalence of self-selection and suggests that characteristics that make firms choose to diversify are negatively correlated with firm value. Hence, firms with a higher probability of diversifying also tend to be discounted.

In Column 5, we additionally include firm fixed effects in the second step of the Heckman selection model. Again the inclusion of firm fixed effects leads to a negative and significant coefficient on the diversification dummy variable while the coefficients on the other variables remain qualitatively similar. To summarize, the results of this section show that by accounting for firm fixed effects, we find a robust and significant conglomerate discount even when the endogeneity of the diversification decision is accounted for in an instrumental variables framework or Heckman's self-selection model.[15]

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[14] Note that the S&P 500 dummy variable is dropped in all regression specifications including firm fixed effects as there is no time-series variation in this variable.

[15] Alternatively, we reestimate the regressions in Table 7 for the excess value measure including the estimate of the market value of debt. As expected, the results are largely consistent with those in Table 7 and, therefore, are not reported in a table.

We check the robustness of the results in this section by using alternative explanatory variables in the first-stage regressions of the instrumental variables regressions (e.g., including past sales growth or omitting several of the variables originally included) and by varying the instruments. We also use an alternative industry definition based on three-digit NAICS codes and repeat the analysis for the asset-based excess value measure (including the book value of debt or the market value of debt based on the artificial credit rating). However, the results change only immaterially (for brevity we do not report them in a table).

6. Changes in Diversification and Excess Value

In this section, we undertake an alternative analysis to investigate the question of causality, i.e., whether firms that diversify are already trading at a discount prior to the diversification, or whether their value decreases as a result of the diversification. Specifically, we investigate whether a change in the degree of diversification is associated with a change in excess value. If diversified firms already trade at a discount before they diversify, this indicates that it is not diversification that causes the discount but that diversification might be a firm's reaction to poor performance. Comment and Jarrell (1995), for example, find that an increase in the degree of diversification is associated with a significant drop in stock returns while an increase in focus is associated with a substantial increase in stock returns. Their results show that a change of 0.1 in the absolute value of a sales-based Herfindahl-Hirshman index is associated with a stock return of about 4%, and that adding or subtracting one business segment is associated with a difference in returns of about 5%.

We begin our analysis by investigating whether diversified firms already trade at a discount before they diversify or whether a discount appears only after the diversification. Panel A of Table 8 reports the mean and median values of the sales-based excess value measure (including the book value of debt) for the year of a change in diversification or focus, the two years before this change, and the year after the change. We consider previously focused firms diversifying in year *t*, diversified firms increasing the number of segments in year *t*, diversified firms decreasing the number of segments in year *t*, and previously diversified firms refocusing in year *t* separately. Most importantly, the results show that previously focused firms that diversify at some point during our sample period do not trade at a discount before diversification. In addition, diversified firms increasing the number of segments trade only at a small discount before diversification. In contrast, focusing firms trade at a very large discount before they decrease the number of segments in which they were active, or become completely focused. This finding suggests that the increase in focus may be due to external pressure (e.g., by active shareholders).

Panel B reports the results of univariate OLS regressions of the change in excess value between years *t* and *t*-1 on a dummy variable, which is set equal to one if a previously focused firm diversifies (Column 1), a diversified firm increases the number of segments (Column 2), a diversified firm decreases the number of segments (Column 3), and a previously diversified firm refocuses (Column 4), respectively. Consistent with the findings of Comment and Jarrell (1995), we find that an increase in focus is positively related to firm value and a decrease in focus (or increase in diversification) is negatively related to firm value. The effect is stronger: (1) For previously diversified firms that become focused than for diversified firms that decrease the number of segments; and (2) for previously focused firms that become diversified than for diversified firms that increase the number of segments.

Another potential concern with our results is that the documented diversification discount is due to conglomerates purchasing discounted target firms rather than diversification itself (e.g., Graham, Lemmon, and Wolf, 2002). We perform two simple tests to control for the effect of mergers on our results. First, we repeat the analysis in Panel A of Table 8 (previously focused firms diversifying) and exclude all observations which are associated with a merger of the company taking place in the same year. The results (reported at the end of Panel A) show that the discount associated with these firms' diversification strategies is even higher as compared to all previously focused firms diversifying. Hence, the discount is unlikely to be mainly caused by the acquisition of discounted targets. Second, we repeat the analysis in Column 1 of Panel B and exclude increases in diversification taking place in years in which the firm undertakes at least one acquisition (results not reported). Again the results remain qualitatively similar (or become even stronger; the coefficient is -0.038 with a *p*-value equal to 0.000) indicating that the conglomerate discount in our sample is not due to the acquisition of discounted targets.

To summarize, the results in this section show that increases in diversification – whether by acquisition or organic growth – are associated with lower firm values. In contrast, increases in focus are accompanied by increases in firm value. Hence, our findings suggest that diversification in fact *causes* the conglomerate discount.

7. Conclusion

This paper contributes to the literature on corporate diversification strategies and the conglomerate discount in three important ways. First, we provide an out-of-sample test of the results of prior literature on the conglomerate discount. Specifically, we consider the post-1997 period where SFAS 131 superseded SFAS 14 in the regulation of segment reporting. Consistent with Berger and Hann (2003), we find a substantial increase in the percentage of diversified firms from roughly 18% to 27% upon the introduction of SFAS 131. In addition, we show that firms revealing their "hidden" degree of diversification experience a substantial drop in excess value. Nevertheless, the results of our multivariate analysis indicate that the conglomerate discount in the post-SFAS 131 is comparable in magnitude to that reported for the pre-1998 period.

Second, we investigate whether diversification has a risk-reducing effect and if such risk effects of diversification affect the conglomerate discount. Mansi and Reeb (2002) argue that the use of the book value of debt may result in an under-estimation of firm value for diversified firms. However, an increase (decrease) in the value of debt can occur only when diversification increases (decreases). If a firm's degree of diversification remains unchanged, bonds are issued at par and their value does not change – unless market rates change, of course. This, however, is not related to the diversification decision of the firms. Hence, we argue that differences in debt value can only emerge when a firm changes the degree of diversification. Even then, they will prevail only for a limited period of time as all subsequent bonds will again be issued at par reflecting the degree of diversification of the firm and the resulting consequences on firm risk. Consequently, we expect the potential effect on the conglomerate discount to be of minor importance and to explain only a small part of the discount, if anything. In fact, when we replace the book value of debt by an estimate of the market value of debt in the calculation of the excess value measure, the discount – although somewhat reduced – remains statistically significant. Only when we rely on the official credit rating provided by S&P in the computation of the excess value measure, the discount disappears in all regression specifications. However, we show that this result is due to the sample selection process resulting from the use of the S&P rating, which is available for 26% of our firm-year observations only. This problem is likely to drive the results of Mansi and Reeb (2002), the only existing study investigating the effect of risk reduction on the conglomerate discount empirically.

Third and most importantly, we investigate the role of the estimation technique on the documented conglomerate discount. Specifically, we use a robust version of the Hausman (1978) specification test to demonstrate the importance of accounting for firm fixed effects when estimating the conglomerate discount. When we include firm fixed effects, the conglomerate discount remains statistically and economically significant in all regression specifications. We additionally account for a potential endogeneity of the diversification decision in an instrumental variables framework or by estimating a Heckman (1979) selection model as proposed by Campa and Kedia (2002) and Villalonga (2004a). When firm fixed effects are neglected, the conglomerate discount in fact disappears as reported in these two studies. However, when firm fixed effects are included, the conglomerate discount is estimated to be statistically significant and in the magnitude of 11% to 21%. Hence, our findings reopen the (re-)search for alternative explanations for the conglomerate discount or the reason for why so many firms diversify or remain diversified if diversification destroys so much value.

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Table 1: Excess Value of Focused and Diversified Firms by Calendar Year

Year	N	Focused (%)	Excess Value	Diversified (%)	Excess Value	Difference
1985	2,510	61.00%	0.0101	39.00%	-0.1697	0.1798 ***
1986	2,573	62.84%	-0.0054	37.16%	-0.1455	0.1401 ***
1987	2,737	66.75%	0.0040	33.25%	-0.0883	0.0924 ***
1988	2,688	68.15%	-0.0032	31.85%	-0.0836	0.0805 ***
1989	2,559	70.14%	0.0067	29.86%	-0.0839	0.0906 ***
1990	2,468	71.39%	0.0051	28.61%	-0.0739	0.0790 ***
1991	2,509	73.34%	-0.0146	26.66%	-0.1073	0.0927 ***
1992	2,755	75.03%	-0.0100	24.97%	-0.0761	0.0662 ***
1993	3,109	77.39%	0.0015	22.61%	-0.1131	0.1146 ***
1994	3,432	78.41%	-0.0017	21.59%	-0.1247	0.1230 ***
1995	3,680	79.70%	-0.0092	20.30%	-0.1045	0.0953 ***
1996	3,990	81.30%	-0.0058	18.70%	-0.1101	0.1043 ***
1997	4,100	81.95%	-0.0133	18.05%	-0.1118	0.0985 ***
1998	3,285	73.49%	-0.0528	26.51%	-0.0749	0.0221
1999	2,995	69.12%	-0.0658	30.88%	-0.1048	0.0391 *
2000	2,920	70.79%	-0.0499	29.21%	-0.0999	0.0500 *
2001	2,759	72.82%	-0.0271	27.18%	-0.1028	0.0757 ***
2002	2,709	73.24%	-0.0090	26.76%	-0.0409	0.0319
2003	2,602	73.21%	0.0033	26.78%	-0.1175	0.1209 ***
2004	2,762	72.27%	0.0161	27.73%	-0.1042	0.1203 ***
2005	2,523	70.91%	-0.0022	29.09%	-0.1398	0.1376 ***

The table reports the number of total observations (*N*), the percentage of focused firms in the sample, the percentage of diversified firms in the sample, and the mean values of the sales-based excess value measure for focused and diversified firms for each sample calendar year. All firms reporting more than one segment (with differing SIC or NAICS) codes are classified as diversified. The equality of means is tested using a standard *t*-test. ***/**/* denotes statistical significance at the 1%/5%/10% level.

Table 2: Conglomerate discount based on multivariate cross-sectional regressions by calendar year (1998-2005)

Year	Coefficient	t-statistic	N	R-squared
Panel A: Excess value based on sales				
1998	-0.097	(-4.162) ***	2,787	0.125
1999	-0.076	(-3.048) ***	2,546	0.110
2000	-0.050	(-1.839) *	2,306	0.108
2001	-0.121	(-4.348) ***	2,334	0.098
2002	-0.098	(-3.623) ***	2,425	0.109
2003	-0.156	(-5.669) ***	2,364	0.082
2004	-0.141	(-5.520) ***	2,482	0.068
2005	-0.165	(-6.184) ***	2,285	0.070
Panel B: Excess value measure based on assets				
1998	-0.104	(-3.665) ***	2,302	0.076
1999	-0.121	(-4.905) ***	2,308	0.056
2000	-0.079	(-2.926) ***	2,207	0.091
2001	-0.097	(-3.770) ***	2,251	0.090
2002	-0.088	(-3.558) ***	2,299	0.083
2003	-0.157	(-6.343) ***	2,245	0.076
2004	-0.157	(-7.041) ***	2,341	0.038
2005	-0.167	(-7.143) ***	2,168	0.057

The table reports coefficient estimates, *t*-statistics, the number of observations, and the R-squared of cross-sectional regressions of the excess value measure based on sales (Panel A) and the excess value measure based on assets (Panel B) on a dummy variable whether the firm is diversified and control variables for each year from 1998 to 2005. The control variables (not reported) include: the natural logarithm of total assets, the ratio of capital expenditures to sales, and the ratio of EBIT to sales. The *t*-values (in parentheses) are based on heteroskedasticity-consistent White (1980) standard errors. ***/**/* denotes statistical significance at the 1%/5%/10% level.

Table 3: What happens between 1997 and 1998?

	1997		1998	
Panel A: Firms changing the number of segments in 1998				
Previously focused firms diversifying	-0.023 (0.000)	[326]	-0.122 (-0.157)	[326]
Diversified firms increasing the number of segments	-0.088 (-0.097)	[52]	-0.145 (-0.248)	[52]
Diversified firms decreasing the number of segments	-0.087 (-0.039)	[59]	-0.068 (-0.011)	[43]
Previously diversified firms becoming focused	-0.199 (-0.306)	[40]	-0.108 (-0.167)	[40]
Panel B: Firms with no changes in the number of segments				
Focused firms: Excess value	0.005 (0.000)	[1,977]	-0.059 (-0.016)	[1,977]
Diversified firms: Excess value	-0.066 (-0.108)	[341]	-0.061 (-0.094)	[341]
Panel C: Firms entering the sample in 1998				
Focused firms: Excess value			0.078 (0.050)	[145]
Diversified firms: Excess value			0.065 (-0.105)	[26]
Diversified firms: Number of Segments			2.269 (2.000)	[26]

Panel A of the table reports the mean and median excess value based on sales (including the book value of debt) of the firms which are in the sample in 1997 and 1998 and change the number of segments between 1997 and 1998 for 1) previously focused firms diversifying, 2) diversified firms increasing the number of segments, 3) diversified firms decreasing the number of segments, and 4) previously diversified firms becoming focused. Panel B reports the mean and median excess value of focused and diversified firms which do not change the number of segments between 1997 and 1998. Panel C reports the mean and median excess value (and the number of segments for diversified firms) for focused and diversified firms entering the sample in 1998. Median values are in parentheses and the number of observations in square brackets.

Table 4: Pooled OLS regressions of the sales-based excess value measure for different sample periods

Dependent Variable: Excess Value Measure Based on Sales (Including the Book Value of Debt)								
Sample Period:	1985-1997	1998-2005	1985-2005	1998-2005	1998-2005	1998-2005	1998-2005	1998-2005
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Constant	-0.497 ***	-0.596 ***	-0.552 ***	-0.680 ***	-0.540 ***	-0.877 ***	-0.596 ***	-0.593 ***
	(-26.786)	(-12.167)	(-19.300)	(-12.095)	(-10.070)	(-32.947)	(-12.141)	(-12.149)
Diversified	-0.132 ***	-0.115 ***	-0.128 ***	-0.118 ***				
	(-20.292)	(-6.715)	(-15.612)	(-5.078)				
Number of Segments					-0.075 ***			
					(-8.845)			
Herfindahl Index (Sales)						0.270 ***		
						(7.563)		
Diversified (unrelated)							-0.096 ***	
							(-5.881)	
Diversified (related)								-0.039
								(-1.126)
ln(Assets)	0.062 ***	0.085 ***	0.075 ***	0.091 ***	0.088 ***	0.086 ***	0.083 ***	0.078 ***
	(17.929)	(11.975)	(13.483)	(7.945)	(12.522)	(12.236)	(11.251)	(10.277)
CAPEX/Sales	0.408 ***	0.246 ***	0.340 ***	0.496 ***	0.243 ***	0.245 ***	0.253 ***	0.264 ***
	(6.887)	(4.945)	(6.349)	(4.704)	(5.059)	(5.032)	(5.138)	(5.420)
EBIT/Sales	0.903 ***	0.132 *	0.370 **	0.395 ***	0.130 *	0.130 *	0.130 *	0.127 *
	(7.103)	(1.958)	(2.498)	(3.365)	(1.920)	(1.907)	(1.906)	(1.807)
Past Sales Growth	0.043 ***	0.062 ***	0.054 ***	0.081 ***	0.062 ***	0.062 ***	0.063 ***	0.065 ***
	(3.716)	(23.552)	(5.766)	(10.769)	(22.566)	(23.203)	(22.404)	(20.763)
R&D/Sales				0.756 ***				
				(7.366)				
N	30,224	19,529	49,753	11,598	19,529	19,515	19,529	19,529
Firms	5,773	5,094	7,746	3,052	5,094	5,093	5,094	5,094
R-squared	0.131	0.090	0.101	0.137	0.092	0.090	0.087	0.083

The table reports estimates from pooled cross-sectional time series regressions of the excess value measure based on sales (including the book value of debt) on different measures of diversification and control variables. The dummy variable measuring unrelated diversification (Column 7) is equal to one if a firm is diversified at the three-digit level and the dummy variable measuring related diversification (Column 8) is equal to one if a firm is diversified at the five- or six-digit level only. Year dummy variables are included in all regressions but are not reported. The *t*-values (in parentheses) are based on Driscoll and Kraay standard errors which are heteroskedasticity-consistent and robust to general forms of cross-sectional and temporal dependence. ***/**/* denotes statistical significance at the 1%/5%/10% level.

Table 5: Pooled OLS regressions of the sales-based excess value measure (1998-2005)

Dependent Variable: Excess Value Measure Based on Sales Including the:	Market Value of Debt (Art. Rating)			MVD (S&P)	Book Value of Debt		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Constant	-0.517***	-0.488***	-0.860***	-0.751***	-0.711***	-0.584***	-1.167***
	(-14.882)	(-16.569)	(-10.178)	(-5.541)	(-4.346)	(-13.032)	(-16.769)
Diversified	-0.088***	0.007	0.012	0.190***	0.092***	0.010	-0.008
	(-6.769)	(0.116)	(0.223)	(3.371)	(4.921)	(0.484)	(-0.330)
ln(Assets)	0.069***	0.071***	0.479***	0.412***	0.409***	0.088***	0.605***
	(13.578)	(15.623)	(11.034)	(8.536)	(8.350)	(13.446)	(26.932)
CAPEX/Sales	0.192***	0.193***	0.095***	0.029	-0.174***	0.246***	0.104***
	(4.016)	(4.027)	(3.037)	(0.661)	(-5.782)	(4.943)	(3.313)
EBIT/Sales	0.159***	0.137***	0.103**	0.455***	0.316***	0.115*	0.104**
	(2.810)	(2.713)	(2.036)	(3.452)	(3.120)	(1.880)	(2.198)
Past Sales Growth	0.054***	0.052***				0.061***	
	(8.123)	(8.105)				(24.356)	
ln(Assets) (1 lag)			-0.145***	-0.240***	-0.184***		-0.155***
			(-5.906)	(-5.157)	(-4.603)		(-5.849)
CAPEX/Sales (1 lag)			0.065*	0.184***	0.130***		0.087***
			(1.672)	(9.309)	(5.535)		(3.719)
EBIT/Sales (1 lag)			-0.013	0.282*	0.214*		-0.058
			(-0.191)	(1.855)	(1.763)		(-1.468)
ln(Assets) (2 lags)			-0.138***	-0.005	-0.094***		-0.165***
			(-3.891)	(-0.197)	(-6.087)		(-8.701)
CAPEX/Sales (2 lags)			0.005	0.129*	0.099**		0.038***
			(0.318)	(1.724)	(2.024)		(3.681)
EBIT/Sales (2 lags)			-0.043	-0.130	-0.098		-0.036
			(-1.285)	(-0.818)	(-0.704)		(-0.885)
ln(Assets) squared			-0.011***	-0.007**	-0.007***		-0.017***
			(-5.056)	(-2.383)	(-3.054)		(-9.719)
Leverage		-0.073*	-0.038	-0.492***	0.054	-0.061**	-0.017
		(-1.758)	(-0.947)	(-7.194)	(1.239)	(-2.319)	(-0.786)
Diversified*Leverage		-0.161	-0.162	-0.382***	-0.233***	-0.222***	-0.170***
		(-1.317)	(-1.503)	(-5.080)	(-8.269)	(-3.790)	(-2.676)
N	16,379	16,379	16,001	4,227	4,227	19,528	19,044
Firms	4,546	4,546	4,440	898	898	5,094	4,982
R-squared	0.060	0.073	0.100	0.121	0.066	0.099	0.143

The table reports estimates from pooled cross-sectional time series regressions of the sales-based excess value measure on a dummy variable whether the firm is diversified and control variables. The excess value measure either includes an estimate of the market value of debt which is derived from an artificial credit rating (Columns 1 to 3), the credit rating provided by S&P (Column 4), or the book value of debt (Columns 5 to 7). Year dummy variables are included in all regressions but are not reported. The *t*-values (in parentheses) are based on Driscoll and Kraay standard errors which are heteroskedasticity-consistent and robust to general forms of cross-sectional and temporal dependence. ***/**/* denotes statistical significance at the 1%/5%/10% level.

Table 6: Fixed effects regressions of the sales-based excess value measure (1998-2005)

Dependent Variable: Excess Value Measure Based on Sales Including the:							
	Market Value of Debt (Art. Rating)			MVD (S&P)	Book Value of Debt		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Constant	-0.578***	-0.568***	-0.633***	-0.504	-1.817***	-1.065***	-1.159***
	(-5.024)	(-4.325)	(-3.748)	(-1.267)	(-11.741)	(-7.814)	(-6.518)
Diversified	-0.112***	-0.067**	-0.047*	0.158*	0.102***	-0.120***	-0.122***
	(-5.514)	(-2.301)	(-1.795)	(1.673)	(2.681)	(-5.095)	(-5.183)
ln(Assets)	0.068***	0.066***	0.513***	0.576***	0.748***	0.150***	0.647***
	(3.744)	(3.295)	(9.257)	(7.679)	(16.978)	(6.755)	(10.628)
CAPEX/Sales	0.302***	0.301***	0.166***	0.172***	0.177***	0.322***	0.197***
	(3.589)	(3.580)	(2.710)	(3.444)	(4.375)	(6.486)	(7.025)
EBIT/Sales	0.163***	0.161***	0.060	0.169*	0.163***	0.154***	0.049*
	(2.649)	(2.641)	(1.213)	(1.752)	(3.169)	(5.088)	(1.819)
Past Sales Growth	0.014***	0.014***				0.018***	
	(4.008)	(3.980)				(4.085)	
ln(Assets) (1 lag)			-0.234***	-0.263***	-0.276***		-0.254***
			(-22.458)	(-6.350)	(-11.069)		(-22.910)
CAPEX/Sales (1 lag)			0.086**	0.113***	0.123***		0.095***
			(2.374)	(2.827)	(4.480)		(6.146)
EBIT/Sales (1 lag)			0.018	0.107*	0.091**		-0.036***
			(0.457)	(1.789)	(1.996)		(-2.796)
ln(Assets) (2 lags)			-0.072***	-0.036	-0.043*		-0.090***
			(-4.570)	(-1.015)	(-1.773)		(-6.386)
CAPEX/Sales (2 lags)			0.015	0.074***	0.096***		0.044***
			(0.732)	(3.185)	(4.109)		(10.605)
EBIT/Sales (2 lags)			0.040***	0.036	0.052**		0.014
			(3.427)	(1.121)	(2.197)		(0.743)
ln(Assets) squared			-0.021***	-0.026***	-0.028***		-0.022***
			(-8.481)	(-7.047)	(-8.440)		(-5.558)
Leverage		-0.001	0.029***	-0.423***	0.152***	0.008	0.043***
		(-0.055)	(2.938)	(-8.622)	(3.879)	(1.288)	(7.676)
Diversified*Leverage		-0.077**	-0.104***	-0.343**	-0.258***	-0.016	-0.002
		(-1.973)	(-2.827)	(-2.329)	(-3.888)	(-0.345)	(-0.029)
N	16,379	16,379	16,001	4,227	4,227	19,528	19,044
Firms	4,546	4,546	4,440	898	898	5,094	4,982
Hausman test	83.09***	1389.50***	114.67***	594.52***	935.79***	803.42***	240.48***
(*p*-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
R-squared	0.030	0.031	0.068	0.086	0.128	0.055	0.111

The table reports estimates from fixed effects regressions of the sales-based excess value measure on a dummy variable whether the firm is diversified and control variables. All regressions include firm- and year-fixed effects. The excess value measure either includes an estimate of the market value of debt which is derived from an artificial credit rating (Columns 1 to 3), the credit rating provided by S&P (Column 4), or the book value of debt (Columns 5 to 7). The *t*-values (in parentheses) are based on Driscoll and Kraay standard errors which are heteroskedasticity-consistent and robust to general forms of cross-sectional and temporal dependence. The reported Hausman test compares the estimates from fixed effects regressions to those from random effects regressions and is robust to heteroskedasticity and general forms of spatial and temporal dependence. ***/**/* denotes statistical significance at the 1%/5%/10% level.

Table 7: Instrumental variables regressions and Heckman's self selection model of the sales-based excess value measure (1998-2005)

Dependent Variable: Excess Value Measure Based on Sales (Including the Book Value of Debt)					
	(1)	(2)	(3)	(4)	(5)
Constant	-1.284 ***	-1.175 ***	-1.154 ***	-1.325 ***	-1.118 ***
	(-17.607)	(-6.578)	(-6.497)	(-4.551)	(-6.116)
Diversified	-0.081	-0.123 *	-0.207 ***	0.022	-0.114 ***
	(-1.418)	(-1.769)	(-2.875)	(0.449)	(-5.779)
ln(Assets)	0.666 ***	0.653 ***	0.650 ***	0.618 ***	0.705 ***
	(23.757)	(10.589)	(10.589)	(6.941)	(9.267)
CAPEX/Sales	0.114 ***	0.191 ***	0.192 ***	0.100	0.210 ***
	(3.916)	(7.131)	(6.824)	(0.415)	(7.008)
EBIT/Sales	0.105 ***	0.045	0.044	0.075	-0.004
	(3.020)	(1.627)	(1.583)	(0.474)	(-0.126)
ln(Assets) (1 lag)	-0.157 ***	-0.260 ***	-0.261 ***	-0.286 ***	-0.279 ***
	(-7.594)	(-24.824)	(-25.454)	(-3.643)	(-25.145)
CAPEX/Sales (1 lag)	0.096 ***	0.094 ***	0.089 ***	-0.043	0.113 ***
	(3.787)	(5.281)	(5.019)	(-0.143)	(6.231)
EBIT/Sales (1 lag)	-0.061 ***	-0.033 **	-0.032 **	-0.146	-0.052 ***
	(-3.352)	(-2.556)	(-2.450)	(-0.625)	(-3.632)
ln(Assets) (2 lags)	-0.168 ***	-0.084 ***	-0.082 ***	-0.107 **	-0.106 ***
	(-11.000)	(-5.900)	(-5.693)	(-2.035)	(-6.303)
CAPEX/Sales (2 lags)	0.041 **	0.043 ***	0.045 ***	0.354 ***	0.063 ***
	(1.992)	(8.925)	(9.009)	(2.899)	(5.094)
EBIT/Sales (2 lags)	-0.037 **	0.015	0.018	0.535 ***	0.033 **
	(-2.075)	(0.925)	(1.035)	(2.871)	(2.272)
ln(Assets) squared	-0.023 ***	-0.022 ***	-0.022 ***	-0.013 **	-0.028 ***
	(-11.213)	(-5.567)	(-5.463)	(-2.324)	(-4.699)
S&P 500 Dummy	0.228 ***			0.174 **	
	(6.296)			(2.112)	
Leverage	-0.020	0.044 ***	0.044 ***	-0.264 ***	-0.068 **
	(-0.515)	(7.876)	(7.904)	(-3.551)	(-2.189)
Lambda				-0.185 **	-0.086 ***
				(-2.204)	(-3.655)
N	18,947	18,947	18,947	16,702	16,702
Firms	4,959	4,959	4,959	4,350	4,350
R-squared	0.142	0.108	0.108	-	-
Firm fixed effects	no	yes	yes	no	yes
Estimation Methodology	IV	IV	IV	Self-select	Self-select

The table reports the estimates from instrumental variables (IV) regressions and Heckman's (1979) self-selection model of the excess value measures based on sales on a dummy variable whether the firm is diversified and control variables. The excess value measure is based on the book value of debt. The first step of the IV regressions in Columns 1 and 2 uses all exogenous variables along with the estimated probability of diversifying whereas the latter is based on a probit regression of the diversification dummy variable on a set of firm-, industry- and time-specific instruments. Column 3 reports the results from a standard two-stage least squares regression. The first step of Heckman's self-selection model consists of a probit regression of a dummy variable whether the firm increases the number of segments on the same firm-, industry- and time-specific instruments as used in the IV regression whereas the sample is restricted to single-segment firms and all diversifying firms. Lambda is the self-selection parameter. Columns 1 and 4 include year dummy variables which are not reported. Columns 2, 3 and 5 include firm- and year-fixed effects and report the within R-squared (Columns 2 and 3). Note that the S&P 500 dummy variable is dropped in Columns 2, 3 and 5 as there is no time-series variation in this variable. The *t*-values in Columns 1, 2, 3, and 5 (in parentheses) are based on Driscoll and Kraay standard errors which are heteroskedasticity-consistent and robust to general forms of cross-sectional and temporal dependence. ***/**/* denotes statistical significance at the 1%/5%/10% level.

Table 8: The valuation effect of changes in diversification and focus

Panel A: Descriptive Statistics

	Excess Value (t+1)	Excess Value	Excess Value (t-1)	Excess Value (t-2)
	(1)	(2)	(3)	(4)
Previously focused firms diversifying				
Mean	-0.080	-0.061	0.007	0.091
Median	(-0.034)	(-0.018)	(0.000)	(0.064)
Obs.	[249]	[255]	[255]	[225]
Diversified firms increasing the number of segments				
Mean	-0.092	-0.001	-0.043	-0.009
Median	(-0.032)	(0.030)	(-0.014)	(-0.005)
Obs.	[119]	[179]	[179]	[155]
Diversified firms decreasing the number of segments				
Mean	-0.136	-0.088	-0.155	-0.106
Median	(-0.134)	(-0.091)	(-0.218)	(-0.075)
Obs.	[122]	[130]	[130]	[102]
Previously diversified firms becoming focused				
Mean	-0.151	-0.008	-0.086	-0.100
Median	(-0.109)	(-0.019)	(-0.129)	(-0.087)
Obs.	[117]	[121]	[121]	[92]
Previously focused firms diversifying (acquisitions excluded)				
Mean	-0.051	-0.096	-0.067	0.072
Median	(-0.021)	(-0.047)	(-0.025)	(0.079)
Obs.	[97]	[100]	[100]	[91]

Panel B: Univariate Regressions of Delta Excess Value

	Focused Firms Diversifying	Diversified Firms Diversifying	Diversified Firms Focusing	Diversified Firms becoming Focused
	(1)	(2)	(3)	(4)
Intercept	-0.056 ***	-0.057 ***	-0.058 ***	-0.058 ***
	(-4.170)	(-4.263)	(-4.207)	(-4.241)
Coefficient	-0.033 ***	-0.003	0.115 ***	0.120 ***
	(-2.955)	(-0.183)	(3.346)	(5.562)
N	20,309	20,309	20,309	20,309
Firms	5,260	5,260	5,260	5,260
R-squared	0.001	0.001	0.001	0.002

Panel A of this table reports mean and median values of the sales-based excess value measure (including the book value of debt) for years *t*, *t*-1, *t*-2, and *t*-3 for previously focused firms diversifying in year *t*, diversified firms increasing the number of segments in year *t*, diversified firms decreasing the number of segments in year *t*, and previously diversified firms refocusing in year *t*. Panel B reports the results of univariate OLS regressions of the change in excess value between years *t* and *t*-1 on a dummy variable, which is equal to one if a previously focused firm diversifies (Column 1), a diversified firm increases the number of segments (Column 2), a diversified firm decreases the number of segments (Column 3), and a previously diversified firm refocuses (Column 4). The numbers in parentheses are *t*-statistics based on Driscoll and Kraay standard errors. ***/**/* denotes statistical significance at the 1%/5%/10% level.