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An alternative test of the trade-off theory of capital structure

Giorgio Canarella[1], Mahmoud Nourayi[2], Michael J. Sullivan[1]

ABSTRACT

The purpose of this paper is to investigate the stochastic behavior of corporate debt ratios utilizing a balanced panel of 2,556 publicly traded U.S. firms during the period 1997-2010. We partition the panel into ten economic sectors and perform panel unit root tests on each sector employing book value and market value measures of debt ratio. First-generation panel unit root tests provide consistent evidence that debt ratios are mean reverting, which supports the trade-off theory. However, these tests rely on the assumption that the debt ratios are cross-sectionally independent, but tests of cross-sectional independence fail to uphold this assumption. Thus, utilizing a second-generation panel unit root test that controls for cross-sectional dependence, we uncover evidence showing that debt ratios are not mean reverting, which contradicts the trade-off hypothesis. We find that the recent macroeconomic developments triggered by the financial crisis and the Great Recession have considerable explanatory power over the dynamics of the debt ratios. In fact, when we exclude the years of the recent global financial crisis, the unit root hypothesis is rejected in one half of the sectors. We interpret these results as indicative that the recent global events may have produced in these sectors a structural change in the underlying data generation process (DGP). Overall, then, we find mixed evidence on the stationarity of debt ratios.

KEY WORDS: panel unit root tests; capital structure theories; cross-sectional dependence; debt ratio

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[1] University of Nevada, Las Vegas - Lee Business School, United States; [2] Loyola Marymount University, United States

Introduction

Since the seminal work of Modigliani and Miller (1958), three main theories have been advanced to explain corporate capital structure: the trade-off theory, the pecking order theory, and the market-timing hypothesis. The trade-off theory is centered on the idea that firms have an "optimal" capital structure that presupposes a target debt ratio and explains this target debt ratio as a trade-off between tax and other benefits against financial distress and other costs that are consequences of the use of debt (Bradley, Jarrell, & Kim, 1984; Graham & Harvey, 2001; Harris & Raviv, 1991; Kraus & Litzenberger, 1973). The pecking order theory, however, postulates that the cost of financing increases with asymmetric information and, therefore, predicts that a firm's debt ratio simply reflects a hierarchy of financing sources whereby internal financing is preferred over debt, and debt is preferred over equity (Myers, 1984; Myers & Majluf, 1984). The market timing theory speculates that capital structure decisions

Correspondence concerning this article should be addressed to: **Mahmoud Nourayi** Loyola Marymount University - Accounting, 1 LMU Drive Hilton 317, Los Angeles, California 90045, United States E-mail: mnourayi@lmu.edu

are driven by firms' attempts to time the equity markets (Baker & Wurgler, 2002). Tests of the trade-off theory attempt to measure the extent and speed of rebalancing a firm's debt ratio towards the presumed target. Much of this research finds evidence supporting the trade-off theory (Almeida & Philippon, 2007; Flannery & Rangan, 2006; Harris & Raviv, 1991; Hovakimian, Opler, & Titman, 2001; Leary & Roberts, 2005). Research that does not support the trade-off theory includes Lemmon, Roberts and Zender (2008), Hovakimian, Kayhan and Titman (2011). Recent surveys of capital structure theories include Baker and Martin (2011) and Frank and Goyal (2007).

Much of the current research investigating issues related to capital structure is methodologically based on structural modeling. That is, it mainly explores the determinants of the observed capital structure. Recently, this research has relied on a variety of econometric techniques, which include, among others, logit and probit models (Bayless & Chaplinsky, 1991; Helwege & Liang, 1996; Hovakimian et al., 2001), the Fama-MacBeth two-step approach (Fama & French, 2002; Flannery & Rangan, 2006; Hovakimian & Li, 2011; Welch, 2004), structural equation models (Chang, Lee, & Lee, 2009; Titman & Wessels, 1988), non-linear methods (Banerjee, Heshmati, & Wihlborg, 2000; Vilasuso & Minkler, 2001), cross-section regressions (Frank & Goyal, 2003; Hanousek & Shamshur, 2011; Rajan & Zingales, 1995), and Kalman filter techniques (Zhao & Susmel, 2008). This type of research has also benefitted from advances in time series and panel data econometrics. For example, the determinants of firm debt ratios have recently been analyzed in a dynamic framework using fixed-effect panel regressions (Flannery & Rangan, 2006; Huang & Ritter, 2009), fractional dependent estimators (Elsas & Florysiak, 2011), generalized method of moments (GMM) methods (Antoniou, Guney, & Paudyal, 2008; Faulkender et al., 2012; Lemmon et al., 2008), and dynamic panel threshold models (Dang, Kim and Shin, 2012; 2014).

A particular concern about these models that has emerged in recent years is that they fail to include an assessment of the stochastic properties of debt ratios and ignore the issue of cross-sectional dependence. The first problem has been discussed at length by Granger and Newbold (1974) and exposes the econometric results to the spurious regression problem when data are non-stationary, i.e., contain unit roots. The second problem is particularly important in dynamic panel regressions. As noted by Phillips and Sul (2003), this substantially complicates the estimation and inference in dynamic panel models. Phillips and Sul (2003) address this problem from a theoretical perspective and propose an approach that is based on a panel version of the median unbiased estimator (Andrews, 1993).

The motivation of this study is twofold. First, unlike the vast bulk of the extant literature that focuses on the determinants of corporate capital structure, we rely on recent developments in the econometrics of non-stationary dynamic panel data. Specifically, we approach the analysis of the trade-off theory by assessing the stochastic properties of corporate debt ratio from the perspective of the panel unit root methodology. If the debt ratio is represented by a stationary process, shocks affecting the series are transitory, and the debt ratio will eventually return to its target level. Thus, evidence of stationarity supports the trade-off theory, as it characterizes the dynamics of capital structure as mean reverting. This situation, in turn, could be interpreted as an indirect signal of industry stability. Conversely, if the debt ratio evolves as a unit root process, shocks affecting the series have permanent effects, shifting the corporate capital structure from one level to another, which contradicts the trade-off theory. Second, we directly address the question of cross-sectional dependence in panel unit root tests. The application of univariate unit root tests, such as the Augmented Dickey-Fuller (Said & Dickey, 1984) and the Phillips-Perron (Phillips & Perron, 1988) tests, is somewhat commonplace in studies employing time series data. In contrast, the use of unit root tests for panel data is more recent (Im, Pesaran, & Shin, 2003; Levin, Lin, & Chu, 2002; Maddala & Wu, 1999). It is by now a generally accepted argument that the commonly used univariate unit root tests lack power in distinguishing the null hypothesis of unit root from stationary alternatives, and utilizing panel data unit root tests is one way of increasing the power of unit root tests (Choi 2001; Im et al., 2003; Levin et al., 2002). Panel unit root tests exploit both the time-series ($t = 1, 2 \ldots T$) and the cross-section ($i = 1, 2 \ldots N$) dimensions of the underlying data, thereby having more power and greater efficiency than univariate time series unit root tests (Baltagi, 2005). The tests share the null hypothesis of unit root, but differ in the alternative. The LLC test, proposed by Levin et al. (2002), tests for the null hypothesis of the unit root against a homogeneous stationary hypothesis, i.e., the autoregressive parameter constrained to be the same across cross-section units, while the IPS test, suggested by Im et al. (2003), and the Fisher type tests developed by Maddala and Wu (1999) and Choi (2001) test for the null hypothesis of unit root against the heterogeneous alternative, i.e., the autoregressive parameter is allowed to vary across cross-section units. Surveys of panel unit root tests include, among others, Banerjee (1999), Breitung and Pesaran (2008), Gutierrez (2006), and Jang and Shin (2005). Unfortunately, however, testing the unit root hypothesis by employing panel data instead of individual time series is not without complications. In particular, the panel unit root literature has noted that in many empirical applications it may be inappropriate to assume that the cross-section units are independent. Observations on firms, industries, regions and countries normally tend to be cross-correlated and serially dependent (Breitung & Pesaran, 2008). Thus, an important problem in panel unit root tests is whether the cross-sections of the panel are independent. On this issue, the panel unit root literature distinguishes between the first-generation tests, which are developed on the assumption of the cross-sectional independence, and the second-generation tests, which account for the dependence that might prevail across the different units in the panel. If the data are cross-sectionally dependent, the panel unit root literature has demonstrated that the first-generation tests can generally be misleading, in the sense that they expose the tests to significant size distortions. That is, the tests tend to reject the null hypothesis of non-stationarity too often (see, for instance, Choi, 2001; Im et al., 2003; Levin et al., 2002; Maddala & Wu, 1999). Moreover, Pesaran (2007) demonstrates that panel unit root tests that do not account for cross-sectional dependence when cross-sectional dependencies are indeed present are seriously biased if the degree of cross-sectional dependence is sufficiently large. To date, only a few studies examine the corporate capital structure employing panel unit root tests. Chang, Liang, Su and Zhu (2010) use quarterly data over the period 1996:Q4-2007:Q3 from a panel of Taiwanese electronic firms and fail to reject the null hypothesis of unit root, except for the subsample of firms with low profitability. Bontempi and Golinelli (2001) utilize annual data from 5,079 Italian firms during the period 1982-1995 and find evidence that favors the trade-off theory. Tasseven and Teker (2009) employ annual data from 42 Turkish firms during the period 2000-2007 and report findings that do not provide support for the trade-off hypothesis. These studies employ first-generation panel unit root tests. Chang, Liang, Su and Zhu (2010) make use of the LLC (Levin et al., 1992) test, the IPS (Im et al., 2003) test, and the Maddala and Wu (Maddala & Wu, 1999) Fisher type tests. Bontempi and Golinelli (2001) apply the IPS test (Im et al., 2003), while Tasseven and Teker (2009) employ the LLC (Levin et al., 1992) and the Maddala and Wu (Maddala & Wu, 1999) Fisher type tests. Thus, all three studies rely upon the assumption of cross-sectional independence.

A large amount of the current research on panel data concentrates on how to address cross-sectional dependence. The second-generation tests, such as the Seemingly Unrelated Regressions Augmented Dickey-Fuller test (SURADF) developed by Breuer, McNown and Wallace (2002), and the Cross-Sectionally Augmented ADF test (CADF) proposed by Pesaran (2007) address explicitly the problem of cross-sectional dependence. The SURADF test is based on a system of augmented Dickey-Fuller (ADF) equations and estimates the autoregressive process by the Seemingly Unrelated Regression Equations (SURE) procedure; i.e., it accounts for cross-sectional dependence by directly incorporating the variance-covariance matrix of the residuals of the equations system in the estimation process. The advantage of this approach is that it allows identification of the cross-sectional units of the panel that contain a unit root (Lau, Baharumshah, & Soon, 2013). The major drawback, however, is that if $N > T$, i.e., the number of cross-section units exceeds the number of time periods, the SURE approach is not feasible. This limitation is also present in the robust version of the non-parametric panel unit root test proposed by Breitung and Das (2005) to account for cross-sectional dependence. In the panel data that we use in our empirical analysis, the number of time periods is significantly less than the number of cross-sections. This fact, in turn, precludes the use of the SURADF test or the Breitung test in our empirical analysis. Instead, we employ a second-generation panel unit root test that allows for cross-sectional dependence developed by Pesaran

(2007), accounting for cross-sectional dependence by imposing a common factor structure. Pesaran (2007) suggests a cross-sectionally augmented Dickey-Fuller (CADF) test where the standard ADF regressions are augmented with cross-sectional averages of lagged levels and first differences of the individual series. The data generating process (DGP) is a simple dynamic linear heterogeneous panel data model. The error term is assumed to have an idiosyncratic component and an unobserved common factor structure accounting for cross-sectional correlation.

There are a variety of reasons why cross-sectional dependence may exist in an industry. Commonly, cross-sectional dependence reflects the fact that firms in the same industry respond to unobserved common stochastic shocks and are linked by unobserved common stochastic trends. Common shocks and common trends spread across all firms in an industry, thus engendering the panel feature of cross-sectional dependence. Monetary and fiscal shocks frequently provide the channels that generate common stochastic shocks. For example, monetary shocks in the supply of money and fiscal shocks in the supply of government debt affect the rate of inflation and the structure of interest rates, which in turn influence the firm's cost of capital and the equilibrium of financial markets, leading to changes in the financial constraints in the corporate sector and alternative representations of the corporate capital structure (Bokpin, 2009; Frank & Goyal, 2009). Furthermore, in a globalized economy, shocks generated in one country are known to cross national borders (Lau, Baharumshah and Soon, 2013). This phenomenon is especially true for oil shocks. The global financial crisis is arguably one of the deepest exogenous shocks that recently affected the corporate sector. The credit supply shock (Dang et al., 2014) originated by the subprime crisis has affected the corporate demand for and supply of funds and, consequently, the capital structure. Common stochastic trends, however, are another source of cross-sectional dependence, as they reflect the presence of corporate variables that tend to move together, i.e., are cointegrated in a VAR system (Granger, 1981). Empirical evidence, for instance, has found that stable relationships exist at the industry level between measures of firm performance, such as sales or profitability, and research and development expenditures (Chan, Lakonishok, & Sougiannis, 2001) and between the market value added of the firm (MVA), an external measure of a firm's performance, and several internal measures, such as earnings per share (EPS), free cash flow per share (FCF), return on equity (ROE), return on assets (ROA), and economic value added per share (EVA) (Bernier & Mouelhi, 2012).

This study contributes to the empirical capital structure literature in several ways. First, as mentioned above, our methodological approach enables us to fill a gap in the existing literature by focusing on an alternative stochastic process that might be more consistent with the long-run behavior of debt ratios. Existing empirical work has focused almost exclusively on the relationships between corporate capital structure and its determinants. While these studies have produced a great deal of evidence on the association between capital structure and its determinants, they have not been able to provide much evidence on the dynamics of debt ratios. Our methodology is based on a panel unit root test that allows for alternative assumptions of cross-sectional dependency for capital structure adjustments. Surveys of panel unit root tests include, among others, Breitung and Pesaran (2008), Banerjee (1999), Gutierrez (2006), and Jang and Shin (2005). Panel unit root tests exploit both the time-series (t = 1, 2…T) and cross-section (i = 1, 2…N) dimensions of the underlying data, thereby having more power and greater efficiency than conventional time series unit root tests (Baltagi, 2005). This type of analysis is not new in corporate finance. Tippett (1990), for example, models financial ratios in terms of stochastic processes, and Tippett and Whittington (1995) and Whittington and Tippett (1999) report empirical evidence that the majority of financial ratios exhibit random-walk behavior. A unit root process imposes no bounds on how a series moves. If the debt ratio really conforms to a random-walk process, then it is unpredictable. A presumption of the trade-off theory is that managers make capital structure decisions based on a target debt ratio and that shocks affecting the debt ratio will prove transitory. This implies that debt ratios are mean reverting towards a target level and follow a stationary dynamic. Conversely, if managers do not make decisions based on a target debt ratio, shocks result in permanent shifts in the debt ratio. In this case, change in the debt ratio evolves as a unit root, non-stationary process, which is consistent with alternative capital structure theories, such as the pecking order or the market timing theories (Baker & Wurgler, 2002; Myers & Majluf, 1984). Non-stationarity of the debt ratio differs from persistence. Persistence involves a slow process of adjustment to an optimal level, while non-stationarity implies that debt ratios fluctuate randomly, driven only by stochastic shocks without a tendency to return to a mean. Therefore, non-stationarity implies that firm debt ratios exhibit a unit root, while persistence suggests that firm debt ratios exhibit a near unit root. It is important to note that the dynamic partial adjustment models currently utilized in the literature are based on assumptions that capital structure adjustments are mean reverting and these adjustments are cross-sectionally independent across firms (Fama & French, 2002; Flannery & Rangan, 2006; Frank & Goyal, 2003; Huang & Ritter, 2009; Leary & Roberts, 2005; Shyam-Sunder & Myers, 1999; Welch, 2004). Evidence on the stochastic properties of the debt ratios also possesses well-defined implications for econometric modeling and forecasting. Failure to reject the unit root hypothesis potentially implies that debt ratios exhibit a long-run cointegrating relationship with other firm-level data, while rejecting the unit root hypothesis implies that debt ratios exhibit only a short-term relationship with other corporate series. Rejecting or not rejecting the unit root hypothesis, in turn, profoundly affects the forecasting process because forecasting based on a mean-reverting process proves quite different from forecasting based on a random walk process.

Second, we control for effects related to the economic sector when analyzing the stochastic properties of debt ratios. We accomplish this by stratifying the data into ten sectors and examining the stochastic properties of debt ratios within each sector. Debt ratios have been found to exhibit significant differences across sectors (Bradley et al., 1984; Lemmon et al., 2008). Graham and Harvey (2001) found that one third of their sample had debt ratios lower than 0.20, and another third had debt ratios higher than 0.40. This stratification is done because of distinct differences in debt ratios across economic sectors, and the extent and speed of reversion of a firm's debt ratio to its target may vary by sector. Firm-level data for each sector are obtained by partitioning a large panel of 2,556 U.S. public companies during the period 1997-2010. Because we partition the sample into sectors, we employ the average debt ratio for the sector as a benchmark. We first examine the evolution of the debt ratios over the entire sample period 1997-2010. The financial literature, however, has recognized that the turbulent and volatile macroeconomic environment created by the recent financial crisis and the resulting Great Recession had severe effects on corporate financial policies (Campello, Graham, & Harvey, 2010; Campello et al., 2011; Duchin, Ozbas, & Sensoy, 2010). Thus, it would seem prudent to evaluate the robustness of the panel unit root results with the events of the global financial crisis and the Great Recession. To account for this problem, we date the financial crisis with the year of the Lehman Brothers bankruptcy. We then construct a "pre-crisis" sub-sample, 1997-2007, and investigate whether this sample reduction has affected our findings. In this respect, our paper adds to the nascent literature that documents the negative impact of the recent financial crisis on corporate debt ratios (Dang et al., 2014).

Third, we measure debt ratios using both book values and market values. Book value and market value debt ratios are conceptually different. Book measures are by definition "backward looking" because of their reliance on accounting data, whereas market values are generally held to be "forward looking". Therefore, differences between the movement of book value and market value debt ratios may be sizeable (Barclay & Morellec, 2006). Rajan and Zingales (1995) and Welch (2004) provide in-depth rationale for analyzing both.

The main findings of our paper can be summarized as follows. First, we find that cross-sectional dependence does matter and substantially affects the outcome of the tests. When we apply conventional, first-generation panel unit root tests that are based on the assumption of cross-sectional independence, we find results that lead to the rejection of the unit root hypothesis. This evidence is consistent with mean reversion of debt ratios and, therefore, supports the trade-off hypothesis. However, to determine if these first-generation tests are appropriate, we utilize diagnostic tests developed by Pesaran (2004) and Frees (1995; 2004). Second, we find strong evidence of substantial cross-sectional dependence within our sample indicating that the assumption of cross-sectional independence is inappropriate. Third, the Pesaran (2007) panel unit root test that allows for cross-sectional dependence consistently yields results supporting the unit root hypothesis, which is inconsistent with debt ratios being mean reverting. This evidence is contradictory to the trade-off hypothesis. Of course, the failure to formally reject a null hypothesis of unit root does not, on its own, rule out the existence of some important structural change. Fourth, in light of this possibility, we find that the recent financial crisis does matter and substantially affects the results of the tests. When the years of the recent financial crisis are excluded from the analysis, the evidence of stationarity re-emerges in one half of the sectors. We interpret these results as providing some indirect evidence that in one half of the sectors the recent global events have caused a structural break in the underlying data generation process (DGP). Thus, overall, our empirical results provide only mixed evidence in favor of the trade-off theory.

The rest of the paper is organized as follows. Section 2 presents a simple dynamic linear autoregressive model of the debt ratios and shows its connections with panel unit root tests. Section 3 describes the sample data and their sources. Section 4 gives a concise outline of the procedures employed in this study and presents the empirical results. Conclusions are presented in Section 5.

Panel Unit Root Tests and the Corporate Debt Ratio

In this section, we outline a dynamic panel model of corporate debt ratios that provides a theoretical background for the application of panel unit root tests. Let $d_{i,t}$ be the debt ratio of firm i, i = 1…N, at time t, t = 1…T. The trade-off hypothesis implies that deviations of the debt ratio, $d_{i,t}$ from the target debt ratio, d_i^* for firm i at time t, are transitory. We assume the target debt ratio is constant over time and firms move towards this target in the long run, considering the trade-off between the marginal costs and benefits of raising funds through issues of debt and equity. Under this hypothesis, the debt ratio $d_{i,t}$ is mean reverting, implying the following stationary stochastic process for $d_{i,t}$

$$d_{it} = d_i^* + \vartheta_{it} \tag{1}$$

where

$$\vartheta_{it} = \sum_{j=1}^{k+1} \beta_{ij} \vartheta_{it-j} + \varepsilon_{i,t} \tag{2}$$

with $\left|\sum_{j=1}^{k+1} \beta_{ij}\right| < 1$ and $\varepsilon_{it}$is a zero-mean white noise process. Equations (1) and (2) jointly imply the following stationary autoregressive process,

$$d_{it} = a_i + \sum_{j=1}^{k+1} \beta_{ij} d_{it-j} + \varepsilon_{it} \tag{3}$$

where $a_i = d_i^* \left(1 - \sum_{j=1}^{k+1} \beta_{ij}\right)$.

Equivalently, equation (3) can be given the augmented Dickey-Fuller (ADF) representation:

$$\Delta d_{it} = a_i + \rho_i d_{it-1} + \sum_{j=1}^{k} \alpha_{ij} \Delta d_{it-j} + \varepsilon_{it} \tag{4}$$

where Δ is the difference operator, and $\rho_i = \sum_{j=1}^{k+1} \beta_{ij} - 1$, $\beta_{ij} = \alpha_{ij} - \alpha_{ij-1}$ for $j = 2,\ldots,k$, and $\beta_{i1} = 1 + \rho_i + \alpha_{i1}$. Solving equation (4) for $\rho_i = 0$ reduces to the unit root process

$$\Delta d_{it} = a_i + \sum_{j=1}^{k} \alpha_{ij} \Delta d_{it-j} + \varepsilon_{it} \tag{5}$$

Equation (5) implies that when there is a shock ε_{it} at time t, the debt ratio changes in the long run by $\left(1 - \sum_{j=1}^{k} \alpha_{ij}\right)^{-1}$. In other words, this suggests the shock has a permanent effect, which is inconsistent with the trade-off hypothesis. Under the null hypothesis, $H_0 : \rho_i = 0$ for all i, the stochastic process describing the debt ratio has a unit root. Under the alternative hypothesis, $H_1 : \rho_i < 0$ for some i, the debt ratio responds to shocks with a mean-reverting process. Therefore, if the empirical results provide evidence of a mean reversion of debt ratios, the trade-off hypothesis is validated. Conversely, if the results provide evidence of a unit root, the debt ratio is not mean reverting, evidence that contradicts the trade-off hypothesis. Finding evidence of a unit root is generally consistent with the pecking order or the market-timing theories.

Data

We employ a panel of annual data on 2,556 publicly traded firms from the U.S. covering all sectors of the economy for the period 1997-2010. Data are obtained from the annual Compustat files, yielding a balanced panel of 35,784 firm-year observations. These sample data include both financially sound firms and those in financial distress to avoid survival bias because the probability of bankruptcy may have a significant impact on a firm's financing decisions. A balanced panel bypasses the potential selection effects that may emerge from specific characteristics of firms entering and leaving the data within the sample period. For this reason, we restrict our period of analysis to the period 1997-2010 and construct a balanced panel of 2,556 firms. A balanced panel is also a requirement of the econometric techniques employed in the analysis. We also stratify the sample into ten economic sectors, following the Compustat economic sector (ECNSEC) classification scheme, and perform panel unit root tests on each sector utilizing our two alternative debt ratio measures. The ten sectors are (the number of firms is reported in parenthesis): 1) Materials (187); 2) Consumer Discretionary (420); 3) Consumer Staples (135); 4) Health Care (361); 5) Energy (181); 6) Financials (261); 7) Industrials (420); 8) Information Technology (449); 9) Telecommunication Services (45); 10) Utilities (97). A summary description of the ten sectors is presented in the Appendix. Firms in the Utilities and Financials sectors are included despite their atypical capital structure, as they are analyzed independently from the rest of the sample. For example, a high debt ratio is normal for financial firms, but the same high debt ratio for non-financial firms may indicate financial distress.

There is no widespread consensus in the literature regarding a single empirical measure of capital structure, in particular, whether the definition of leverage should utilize book values or market values. A discussion concerning the different measures of leverage can be found in Titman and Wessels (1988) and Rajan and Zingales (1995). Myers (1977) and Fama and French (2002) favor the use of book values, while Welch (2004) advocates the use of market values. Drobetz, Pensa and Wanzenried (2007) discuss the advantages and disadvantages of each measure. We follow Rajan and Zingales (1995) and define leverage as the ratio of financial debt to debt plus equity. We include short-term debt in the definition of the debt ratio as its omission may lead to an understatement of financial distress risk. We consider both the book and market values of equity because it is highly possible that some firms operate within a book value framework rather than a market value framework, and vice versa. The book value of the debt ratio BDR_{it} of firm i at time t is defined as follows:

$$BDR_{it} = \frac{LTD_{it} + STD_{it}}{LTD_{it} + STD_{it} + BVE_{it}} \qquad (6)$$

where LTD_{it} is the book value of long-term debt (Compustat annual data item 9), STD_{it} is the book value of short-term debt (Compustat annual data item 34), BVE_{it} equals the book value of equity computed as the difference between the value of total assets (Compustat annual data item number 6), the sum of LTD_{it} is the book value of long-term debt (Compustat annual data item 9) and STD_{it} is the book value of short-term debt (Compustat annual data item 34). Alternatively, the denominator in equation (6) equals TA_{it}, total assets (Compustat annual data item 6). Similarly, the market value of the debt ratio MDR_{it} of firm i at time t is defined as follows:

$$MDR_{ir} = \frac{LTD_{it} + STD_{it}}{LTD_{it} + STD_{it} + MVE_{it}} \qquad (7)$$

where MVE_{it} is the market value of equity, computed as $\eta_{it}P_{it}$ where η_{it} is the number of shares outstanding (Compusta*t* annual data item 54) and P_{it} denotes the stock price (Compustat annual data item 199).

In Table 1 we report the pooled mean, standard deviation, and median of the book and market value debt ratios for each of the ten economic sectors. Because we utilize ratios of variables, a transformation of the variables to constant prices is not necessary. Our data reveals that there are considerable differences in these two measures of debt ratio across most economic sectors. The average book value debt ratios are higher than the corresponding market value debt ratios with the exception of Financials and Utilities and exhibit a higher standard deviation than the market value debt ratios with the exception of Utilities. In the case of Utilities, however, the standard deviation of the book value debt ratio (0.153) is not significantly different from the standard deviation of the market value debt

Table 1. Descriptive statistics

Economic Sector	Book value debt ratio			Market value debt ratio		
	Mean	Std. dev.	Median	Mean	Std. dev.	Median
Materials	0.679	10.473	0.236	0.261	0.23	0.213
Consumer Discretionary	0.374	2.835	0.227	0.266	0.258	0.193
Consumer Staples	0.316	0.906	0.231	0.232	0.229	0.162
Health Care	0.371	2.056	0.102	0.132	0.196	0.047
Energy	0.262	0.497	0.209	0.231	0.228	0.171
Financials	0.329	0.472	0.254	0.361	0.283	0.324
Industrials	0.274	1.386	0.202	0.239	0.229	0.178
Information Technology	0.391	5.758	0.059	0.124	0.183	0.036
Telecommunication Services	0.594	3.107	0.356	0.328	0.232	0.297
Utilities	0.378	0.153	0.363	0.439	0.155	0.436

Note: The table reports the mean, standard deviation and median of the book value and market value debt ratio for each of the ten economic sectors. All data are from the Compustat database. Debt ratios are calculated for 2,556 publicly traded U.S. corporations from 1997 to 2010. The book value of the debt ratio BDR_{it} of firm i at time t is defined as $BDR_{it} = \frac{LTD_{it} + STD_{it}}{LTD_{it} + STD_{it} + BVE_{it}}$, where LTD_{it} is the book value of long-term debt, STD_{it} is the book value of short-term debt, and BVE_{it} equals the book value of equity. The market value of the debt ratio MDR_{it} of firm i at time t is defined as $MDR_{ir} = \frac{LTD_{it} + STD_{it}}{LTD_{it} + STD_{it} + MVE_{it}}$, where MVE_{it} is the market value of equity, computed as $\eta_{it}P_{it}$ where η_{it} is the number of shares outstanding and P_{it} denotes the stock price. Sectors are defined by the Compustat economic sector (ECNSEC) classification system. The ten sectors (number of firms in parenthesis; Compustat economic sector code in brackets) are: 1) Materials (187) [1000]; 2) Consumer Discretionary (420) [2000]; 3) Consumer Staples (135) [3000]; 4) Health Care (361) [3500]; 5) Energy (181) [4000]; 6) Financials (261) [5000]; 7) Industrials (420) [6000]; 8) Information Technology (449) [8000]; 9) Telecommunication Services (45) [8600]; 10) Utilities (97) [9000].

ratio (0.155). These findings are contrary to the argument that book debt ratios are less subject to uncontrollable firm factors, such as market price variability. We also find that the empirical distributions of each measure of debt ratio are generally non-symmetric. In each case, the mean is greater than the median, implying that the distribution is positively skewed (longer tails to the right).

A number of firms demonstrating extremely leveraged positions are found in the sectors Materials, Consumer Discretionary, Industrials, Information Technology, and Telecommunication Services. Extreme leverage, defined as those debt ratios beyond two standard deviations, are left in the analysis to avoid use of an arbitrarily selection criterion. In cases where there was evidence of extreme values, the analysis also conducted "winsorizing" of the top and bottom 5% of the data. This approach was used to eliminate any unexpected effects of outliers. There was no meaningful effect on the results. We do not report the winsorized results, but these are available upon request. Table 1 reveals the potential problem of aggregating all observations as opposed to stratifying them by sector. The four largest sectors in terms of the number of observations (Information Technology, Consumer Discretionary, Materials, and Health Care) account for approximately 65% of the total number of observations. The remaining six sectors may have little, if any, impact on parameter estimates and test results in an aggregated sample. Under these circumstances, pooled regressions are likely to primarily reflect the behavior of only a few large sectors.

Empirical Results

We present the empirical evidence in three stages. First, we perform the panel data statistical analysis utilizing Fisher type tests (Choi, 2001; Maddala & Wu, 1999). These tests are nonparametric and have the advantage of allowing for as much heterogeneity across units as possible. They belong to the first generation of panel unit root tests, which include among others, Levin et al., (2002), Im et al., (2003), Harris and Tzavalis (1999). The findings of these tests uniformly favor the trade-off hypothesis. However, the concern is that these tests are not robust in the presence of cross-sectional dependence. In other words, these first-generation tests employ a methodology that incorporates the often implausible assumption of cross-sectional independence and fail to discriminate between stationarity with cross-sectional independence and non-stationarity with cross-sectional dependence. The power of the conventional panel unit root tests is weakened by the presence of cross-sectional dependence. Therefore, we next test this assumption of cross-sectional independence utilizing the approaches suggested by Pesaran (2004) and Frees (2004). We find exhaustive evidence indicating the presence of heterogeneous cross-sectional dependencies among the time series, which calls for an alternative test methodology. Consequently, we utilize a second-generation panel unit root test that accounts for cross-sectional dependence based on the methodology of Pesaran (2007). We conclude this section with an assessment of the impact of the recent financial crisis on the results obtained using the full sample.

Results of the panel unit root tests under the assumption of cross-sectional independence

We first implement the Fisher type (Fisher, 1932) unit root tests developed by Maddala and Wu (1999) and Choi (2001). The tests allow for heterogeneity in the parameter estimates and combine the evidence on the unit root hypothesis from the individual unit root tests performed on each cross section unit of the panel. The null hypothesis in the Fisher type tests (and in the *IPS* tests) is the hypothesis of unit root, and the alternative is that of stationarity. This hypothesis is reversed in the Hadri test, in which the null hypothesis is one of stationarity and the alternative is the unit root (Hardi, 2000).

From a meta-analysis perspective, these tests combine the p-values of N univariate independent unit root tests using the inverse chi-square, inverse normal, and inverse logit transformations and are more powerful than the test proposed by Im et al. (2003). The Maddala and Wu (1999) test statistic is defined as follows:

$$P = -2\sum_{i=1}^{N}\ln(p_i) \tag{8}$$

where p_i is the p-value of the test statistic in cross-section unit i. P is the inverse chi-square test, distributed chi-square with $2N$ degrees of freedom under the null hypothesis of a unit root in each cross section. Additionally, Choi (2001) proposes the Z, L^*, and P_m tests, based on the combination of individual p-values. Z is the inverse normal test, distributed as a standard normal $N(0,1)$,

$$Z = \frac{1}{\sqrt{N}}\sum_{i=1}^{N}\Phi^{-1}(p_i) \tag{9}$$

where Φ is the standard normal cumulative distribution function. The L^* test is represented as

$$L^* = \sqrt{k}L \tag{10}$$

where $k = \frac{3(5N+4)}{\pi^2 N(5N+2)}$ and L^* has a t distribution with $5N+4$ degrees of freedom. L is referred to as the inverse logit test and has the logistic distribution with mean 0 and variance $\frac{\pi^2}{3}$.

$$L = \sum_{i=1}^{N}\ln\left(\frac{p_i}{1-p_i}\right) \tag{11}$$

The P_m test is a modified version of the Maddala and Wu's (1999) P test applied to large panels because in the limit the P test statistic has a degenerate distribution.

$$P_m = \frac{1}{2\sqrt{N}}\sum_{i=1}^{N}(-2\ln p_i - 2) \tag{12}$$

where $E[-2p_i] = 2$ and $\text{var}[-2\ln p_i] = 4$, which converges to a standard normal distribution.

We report the results of the Maddala and Wu (1999) P test and Choi (2001) Z test applied to the book value and market value debt ratios in Tables 2 and 3. For robustness reasons, we report the test statistics computed using the p-values from both the Augmented Dickey-Fuller (ADF) and the Phillips-

Table 2. Fisher type panel unit root test results for book value debt ratios

Economic Sector	Fisher type ADF tests		Fisher type PP tests	
	P	Z	P	Z
Materials	944.854*	-19.058*	2377.268*	-39.637*
	(0.000)	(0.000)	(0.000)	(0.000)
Consumer Discretionary	2492.886*	-31.507*	1877.109*	-11.398*
	(0.000)	(0.000)	(0.000)	(0.000)
Consumer Staples	663.841*	-14.658*	362.974*	-1.522***
	(0.000)	(0.000)	(0.0001)	(0.064)
Health Care	1499.432*	-18.557*	2048.728*	-15.949*
	(0.000)	(0.000)	(0.000)	(0.000)
Energy	1236.854*	-23.972*	787.791*	-12.784*
	(0.000)	(0.000)	(0.000)	(0.000)
Financials	1634.913*	-25.631*	1589.257*	-17.771*
	(0.000)	(0.000)	(0.000)	(0.000)
Industrials	2842.283*	-35.836*	2107.669*	-22.300*
	(0.000)	(0.000)	(0.000)	(0.000)
Information Technology	1395.86*	-15.29*	2571.826*	-16.880*
	(0.000)	(0.000)	(0.000)	(0.000)
Telecommunication Services	258.731*	-9.754*	159.170*	-3.407*
	(0.000)	(0.000)	(0.000)	(0.0003)
Utilities	575.050*	-14.942*	305.714*	-4.544*
	(0.000)	(0.000)	(0.000)	(0.000)

Note: See Table 1. The table reports the values of the P test statistic and Z test statistic for book value debt ratios for both the Augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests. The P statistics is computed as $P = -2\sum_{i=1}^{N}\ln(p_i)$ where p_i is the p-value of the test statistic in the cross-sectional unit i. P is the inverse chi-square test and is distributed as a chi-square distribution with 2N degrees of freedom under the null hypothesis of a unit root in each cross-section. The Z test statistic is the inverse normal statistic, is computed as $Z = \frac{1}{\sqrt{N}}\sum_{i=1}^{N}\Phi^{-1}(p_i)$ and is distributed as a standard normal $N(0,1)$. The p-values appear in parenthesis under the test statistics. * denotes significance at the 1% level; *** denotes significance at the 10% level.

Perron (PP) tests conducted on each panel. In the Fisher type *ADF* unit root tests, we rely on the Akaike information criterion (*AIC*) as a lag selection procedure and allow for a maximum lag of 3. The AIC is defined as $-2(LL/T)+2(k/T)$, where LL is the log of the likelihood function with k parameters estimated using T observations. In the Fisher type PP unit root test, the spectral regressions employ the Bartlett kernel in conjunction with the Newey-West bandwidth selection. For economy of space, we do not report the results of the L^* and P_m tests because they are uniformly consistent with the results of the P and Z tests. We do not include a time trend because a time trend is not consistent with a long-run positive, non-accelerating target debt ratio. However, we do include an intercept because the average debt ratio is nonzero. We perform all tests using the "demeaned" version (i.e., we subtract the cross-sectional means from observed data to reduce the degree of contemporaneous correlation) and, in the Fisher type *ADF* tests, we include one lag (to account for serial correlation) chosen by *AIC*. Subtracting the cross-sectional means from the observed data is a strategy suggested by Levin et al. (2002) and Im et al. (2003) to address cases where disturbances may be correlated across firms.

Table 3. Fisher type panel unit root test results for market value debt ratios

Economic Sector	Fisher type ADF tests		Fisher type PP tests	
	P	Z	P	Z
Materials	1139.09*	-21.21*	624.99*	-7.79*
	(0.000)	(0.000)	(0.000)	(0.000)
Consumer Discretionary	2867.46*	-34.82*	1576.89*	-12.14*
	(0.000)	(0.000)	(0.000)	(0.000)
Consumer Staples	855.64*	-18.99*	488.37*	-5.94*
	(0.000)	(0.000)	(0.000)	(0.000)
Health Care	2396.06*	-31.87*	1591.91*	-15.8*
	(0.000)	(0.000)	(0.000)	(0.000)
Energy	1289.08*	-24.36*	682.06*	-9.82*
	(0.000)	(0.000)	(0.000)	(0.000)
Financials	1636.47*	-25.71*	972.49*	-8.85*
	(0.000)	(0.000)	(0.000)	(0.000)
Industrials	2744.67*	-33.88*	1365.55*	-10.22*
	(0.000)	(0.000)	(0.000)	(0.000)
Information Technology	2991.18*	-36.08*	1964.82*	-17.45*
	(0.000)	(0.000)	(0.000)	(0.000)
Telecommunication Services	294.81*	-11.41*	165.19*	-4.06*
	(0.000)	(0.000)	(0.000)	(0.000)
Utilities	630.88*	-16.25*	324.04*	-4.99*
	(0.000)	(0.000)	(0.000)	(0.000)

Note: See Table 1. The table reports the values of the P test statistic and Z test statistic for market value debt ratios for both the Augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests. The P statistics is defined as $P = -2\sum_{i=1}^{N} \ln(p_i)$ where p_i is the p-value of the test statistic in the cross-sectional unit i. P is the inverse chi-square test and is distributed as a chi-square distribution with 2N degrees of freedom under the null hypothesis of a unit root in each cross-section. The Z test statistic is the inverse normal statistic and is distributed as a standard normal $N(0,1)$, $Z = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \Phi^{-1}(p_i)$. The p-values appear in parenthesis under the test statistics. * denotes significance at the 1% level.

The Fisher type meta-statistics strongly indicate that the unit root hypothesis should be rejected in all cases at any conventional significance level. This finding indicates that the behavior of equation (4) in section 2 is consistent with firms borrowing to gradually adjust toward their target debt ratios. Random shocks have only a transitory effect on the debt ratio. This evidence offers support for the trade-off hypothesis and are consistent with those of Harris and Raviv (1991) and Almeida and Philippon (2007), among others. For robustness, we have also computed the tests proposed by Im et al. (2003), Harris and Tzavalis (1999). The findings uniformly confirm the results presented in Tables 2-3. In all economic sectors, the null hypothesis of a panel unit root is rejected rather strongly. To save space, these results are not presented, but are available upon request.

However, these first-generation Maddala and Wu (1999) and Choi (2001) tests are only valid under the assumption of cross-sectional independence, where the error terms are assumed to be independent across individual cross-sections. A weakness of the "demeaning" transformation to overcome the problem of cross-sectional dependence is the implicit assumption that cross-sectional dependence is homogeneous; i.e., cross-sectional dependence is driven by a common factor that has a homogeneous effect on all firms in the industry, regardless of their size. This assumption is highly unrealistic for most practical settings because it ignores the heterogeneous impact of short-run co-movements (common cycles) and long-run co-movements (common trends) on the dynamics of firms within the same industry (O'Connell, 1998). The presence of heterogeneous cross-sectional dependencies undermines the power of the Maddala and Wu (1999) and Choi (2001) tests, leads to false rejections of the null hypothesis of the unit root, and may produce evidence of stationarity when the data are non-stationary. In the next section, we address this issue by testing for cross-sectional dependence using the diagnostic tests proposed by Pesaran (2004) and Frees (2004).

Results of the tests for cross-sectional dependence

Pesaran (2004) proposes a general test for cross-sectional dependence referred to as the CD test. As demonstrated by Pesaran (2004), the CD test applies to a large variety of panel data models. This includes stationary and non-stationary dynamic heterogeneous panel models having a small T (years) and a large N (firms), which is the case for the sample panel data employed in this study. The CD test applies to both balanced and unbalanced panels, is robust to parameter heterogeneity and structural breaks in the slope coefficients and error variance, and performs well in terms of size and power. Under the null hypothesis, the covariance matrix of the residuals is diagonal, i.e., $H_0: \rho_{ij} = \rho_{js} = \mathrm{corr}(\varepsilon_{it}, \varepsilon_{jt}) = 0$ for $i \neq j$, and ε_{it} is independent and identically distributed over time periods and across cross-sectional units. Under the alternative hypothesis $H_1: \rho_{ij} = \rho_{js} \neq 0$ for some $i \neq j$, ε_{it} is correlated across cross-sections but uncorrelated over time. Under the null hypothesis of cross-sectional independence, the CD test statistic is distributed as a standard normal for a sufficiently large N. The CD test averages the pair-wise correlation coefficients of the residuals obtained from the individual Augmented Dickey-Fuller (ADF) regression equations. We compute the CD test statistic for a balanced panel as follows:

$$\mathrm{CD} = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\rho(\hat{\varepsilon}_i, \hat{\varepsilon}_j)\right) \tag{13}$$

where
$$\rho(\hat{\varepsilon}_i, \hat{\varepsilon}_j) = \frac{\sum_{t=1}^{T}\hat{\varepsilon}_{it}\hat{\varepsilon}_{jt}}{\left(\sum_{t=1}^{T}\hat{\varepsilon}_{it}^2\right)^{1/2}\left(\sum_{t=1}^{T}\hat{\varepsilon}_{jt}^2\right)^{1/2}},$$

and where $\hat{\varepsilon}_{it}$ and $\hat{\varepsilon}_{jt}$ are estimated residuals from the Augmented Dickey-Fuller (ADF) regression equations. Under the null hypothesis of cross-sectional independence, the CD test statistic converges asymptotically to the standardized normal distribution. A possible drawback of the CD test is that it involves the sum of the pair-wise correlation coefficients of the residual matrix, rather than the sum of the squared correlations. This situation implies that the test is likely to miss cases of cross-sectional dependence where the signs of the correlations are alternating (for example, where there are large positive and large negative correlations in the residuals) and cancel one another out in the averaging process. This outcome, however, is not the case in our findings. We computed the average absolute value of the off-diagonal elements of the cross-sectional correlation matrix of residuals. This approach can help identify cases of cross-sectional dependence where the sign of the correlations alternates. The average absolute correlation of the off-diagonal elements for the cross-sectional correlation matrix of residuals ranges from 0.273 (Consumer Staples) to 0.833 (Materials) for the book value debt ratio, and from 0.267 (Health care) to 0.326 (Telecommunication Services) for the market value debt ratio. These estimates further reinforce strong evidence of cross-sectional dependence.

Frees (1995, 2004) proposes a statistic that is not subject to this shortcoming. The statistic is based on the sum of the squared correlation coefficients and is given by:

$$\mathrm{R}^2_{\mathrm{AVE}} = \frac{2}{N(N-1)}\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\rho^2(\hat{\varepsilon}_i, \hat{\varepsilon}_j) \tag{16}$$

where $\hat{\varepsilon}_i$ and $\hat{\varepsilon}_j$ are the residuals obtained from the same models estimated for the *CD* test. Frees (1995, 2004) demonstrates that a function of R^2_{AVE} follows a joint distribution of two independent chi-square variables, i.e.

$$\mathrm{C}_{\mathrm{AVE}} = N(R^2_{AVE} - (T-1)^{-1}) \xrightarrow{d} Q \tag{17}$$

Table 4. Results of the CD test for cross sectional independence for both book value and market value debt ratios

Economic Sector	Book value debt ratio	Market value debt ratio
Materials	353.898*	56.778*
	(0.000)	(0.000)
Consumer Discretionary	483.151*	113.979*
	(0.000)	(0.000)
Consumer Staples	12.596*	23.086*
	(0.000)	(0.000)
Health Care	97.021*	47.057*
	(0.000)	(0.000)
Energy	38.445*	71.933*
	(0.000)	(0.000)
Financials	11.677*	85.796*
	(0.000)	(0.000)
Industrials	46.719*	100.314*
	(0.000)	(0.000)
Information Technology	482.925*	74.251*
	(0.000)	(0.000)
Telecommunication Services	36.472*	27.434*
	(0.000)	(0.000)
Utilities	29.175*	52.284*
	(0.000)	(0.000)

Note: See Table 1. The table reports the CD test statistic for a balanced panel computed as $CD = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\rho(\hat{\varepsilon}_i,\hat{\varepsilon}_j)\right)$ where $\rho(\hat{\varepsilon}_i,\hat{\varepsilon}_j) = \frac{\sum_{t=1}^{T}\hat{\varepsilon}_{it}\hat{\varepsilon}_{jt}}{\left(\sum_{t=1}^{T}\hat{\varepsilon}_{it}^2\right)^{1/2}\left(\sum_{t=1}^{T}\hat{\varepsilon}_{jt}^2\right)^{1/2}}$, and where $\hat{\varepsilon}_{it}$ and $\hat{\varepsilon}_{jt}$ are estimated residuals from the Augmented Dickey-Fuller (ADF) regression equations. * denotes significance at the 1% level.

where

$$Q = a(T)[\chi^2_{1,T-1} - (T-1)] + b(T)[\chi^2_{2,T(T-3)/2} - T(T-3)] \qquad (18)$$

and where $\chi^2_{1,T-1}$ and $\chi^2_{2,T(T-3)/2}$ are independent chi-square random variables with $T-1$ and $T(T-3)/2$ degrees of freedom, respectively, and where $a(T) = 4(T+2)/\left(5(T-1)^2(T+1)\right)$ and $b(T) = 2(5T+6)/\left(5T(T-1)^2(T+1)\right)$. The null hypothesis is rejected if $R^2_{AVE} > (T-1)^{-1} + Q_q/N$ where Q_q is the appropriate quintile of the Q distribution.

We report the findings of the two diagnostic tests in Tables 4 and 5. The outcomes of these tests clearly indicate the presence of cross-sectional dependence in both the book value and market value debt ratios. The tests strongly reject the null hypothesis of cross-sectional independence at any conventional significance level. This situation casts doubt on the statistical evidence in favor of stationarity by the Fisher type tests. In addition, the estimates of the residuals correlation coefficients present a wide range of variability, suggesting that residual correlation is heterogeneous rather than homogeneous. For economy of space, the matrices of the estimates of residual correlation coefficients are not reported, but are available on request.

To summarize, the rejection of the null hypothesis of cross-sectional independence implies that tests for the presence of a unit root in book value and market value debt ratios should take this dependence into account to produce unbiased and reliable test statistics. These findings call into question any conclusions drawn from the Fisher type tests. The next section ad-

Table 5. Results of the C_{AVE} test for cross-sectional independence for both book value and market value debt ratios

Economic Sector	Book value debt ratio	Market value debt ratio
Materials	112.921*	5.474*
Consumer Discretionary	123.255*	10.604*
Consumer Staples	0.896*	3.174*
Health Care	15.716*	7.171*
Energy	9.888*	9.075*
Financials	8.565*	8.754*
Industrials	37.942*	13.599*
Information Technology	108.757*	16.917*
Telecommunications Services	6.438*	2.951*
Utilities	4.899*	5.113*

Note: See Table 1. CAVE is a second cross-sectional dependence test. The CAVE test statistic for a balanced panel is computed as $C_{AVE} = N(R^2_{AVE} - (T-1)^{-1}) \xrightarrow{d} Q$ where $\hat{\varepsilon}_i$ and $\hat{\varepsilon}_j$ are the residuals obtained from the same Augmented Dickey-Fuller (ADF) regressions estimated for the CD test. R^2_{AVE} follows a joint distribution of two independent chi-square variables, i.e. $C_{AVE} = N(R^2_{AVE} - (T-1)^{-1}) \xrightarrow{d} Q$ where $Q = a(T)[\chi^2_{1,T-1} - (T-1)] + b(T)[\chi^2_{2,T(T-3)/2} - T(T-3)]$, $\chi^2_{1,T-1}$ and $\chi^2_{2,T(T-3)/2}$ are independent chi-square random variables with $T-1$ and $T(T-3)/2$ degrees of freedom, respectively, $a(T) = 4(T+2)/\left(5(T-1)^2(T+1)\right)$, and $b(T) = 2(5T+6)/\left(5T(T-1)^2(T+1)\right)$. The null hypothesis is rejected if $R^2_{AVE} > (T-1)^{-1} + Q_q/N$ where Q_q is the appropriate quintile of the Q distribution. The p-values are not available. The 1%, 5%, and 10% critical values for the Q test statistic for T=14 are 0.360, 0.184, and 0.184, and denoted *, **, and ***, respectively.

dresses this issue by applying the test developed by Pesaran (2007) that does not require cross-sectional independence.

Results from the panel unit root tests under the assumption of cross-sectional dependence

In this sub-section, we investigate the stationarity property of the two measures of debt ratio by applying the panel unit root test developed by Pesaran (2007). The test assumes that cross-sectional dependence is present in the data in the form of a single unobservable common factor. The test expands on the Im et al. (2003) panel unit root test by augmenting the ADF regression with the cross-sectional averages of lagged level, and contemporaneous and lagged cross-sectional averages of the first differences of the individual series. The test is a two-step procedure. First, Pesaran (2007) proposes a test on the t-ratio of the OLS estimate of β_i in the following cross-sectionally augmented ADF (CADF) regressions

$$\Delta d_{it} = \alpha_i + \beta_i d_{it-1} + \delta_i \bar{d}_{t-1} + \sum_{i=1}^{p} \lambda_{ij} \Delta d_{it-j} + \vartheta \Delta \bar{d}_t + \sum_{i=1}^{p} \vartheta_{ij} \Delta \bar{d}_{t-j} + \varepsilon_{it} \quad (19)$$

where d_{it} represents the debt ratio of firm i at time t, ε_{it} denotes the regression error term, p is the lag order of the model, and $\bar{d}_t = \frac{1}{N}\sum_{i=1}^{N} d_{it}$, $\Delta\bar{d}_t = \sum_{i=1}^{N} \Delta d_{it}$ are cross-sectional averages, intended as a proxy for the unobserved common factor. The lagged terms $\Delta\bar{d}_t$ and Δd_{it} act to filters out contemporaneous correlation among d_{it}. The null hypothesis of the test can be expressed as $H_0: \beta_i = 0$ for all i compared to the alternative hypothesis $H_1: \beta_i < 0$ for some i. Then, consistent with Im et al. (2003), Pesaran (2007) proposes a cross-sectional augmented version of the IPS test, which is a simple average of the individual CADF tests.

$$\text{CIPS}(N,T) = N^{-1} \sum_{i=1}^{N} t_i(N,T) \quad (20)$$

where $t_i(N,T)$ is the cross-sectionally augmented Dickey-Fuller statistic for the i-th cross-sectional unit, given by the t-ratio of the coefficient of d_{it-1} in the *CADF* regression defined by equation (19). Under the null of non-stationarity, the CIPS statistic has a non-standard distribution even for large *N*, but has good size and power properties even when *T* and *N* are relatively small. The critical values for 1%, 5% and 10%, however, are tabulated in Pesaran (2007). In addition, Pesaran (2007) constructs a truncated version of the CIPS, denoted as CIPS^*, to avoid the problem of an extreme statistic in cases when *T* is small, and to ensure the existence of the first and second moments of $t_i(N,T)$. The truncated test is given by

$$\mathrm{CIPS}^*(N,T) = N^{-1}\sum_{i=1}^{N} t_i^*(N,T) \tag{21}$$

where

$$t_i^*(N,T) = \begin{cases} t_i(N,T) & -K_1 < t_i(N,T) < K_2 \\ -K_1 & t_i(N,T) \le K_1 \\ K_2 & t_i(N,T) \ge K_2 \end{cases} \tag{22}$$

The parameters K_1 and K_2 are positive constants based on Monte Carlo simulations (K_1 = 6.19 and K_2 = 2.61). The limiting distribution of the CIPS and CIPS^* statistics are non-standard, even in cases of a large *N*. The critical values are tabulated in Pesaran (2007) for various combinations of *N* and *T*. In Table 6, we report the results of the CIPS^* tests. The results of the CIPS tests are similar and are therefore not reported. These findings are in sharp contrast to the first-generation panel unit root tests presented in Tables 2-3. While Tables 2-3 provide evidence of stationarity of debt ratios, the evidence presented in Table 6 demonstrates the opposite. After controlling for heterogeneous cross-sectional dependence, the evidence reveals a non-stationary (unit root) debt ratio process in the vast majority of the sectors. Therefore, we cannot reject the null hypothesis of a unit root in all ten sectors for the market value debt ratios, and in nine out of ten sectors for book value debt ratios. This failure to reject the stationarity of debt ratios is consistent with hypotheses that do not envisage the existence of a target debt ratio and an adjustment process toward it. It indicates that borrowing is not driven by an attempt to move toward a target capital structure, but instead indicates that borrowing is driven by a need for external funds that is consistent with the pecking order and market-timing hypotheses. Furthermore, the acceptance of the unit root indicates that random shocks have permanent effects on a firm's capital structure. Debt ratios behave as a stochastic process driven year-after-year by external shocks that affect firms. To demonstrate the robustness of our findings, we checked whether our results are sensitive to our measures of debt ratio. We repeated the panel unit root tests presented in Tables 2-6 utilizing two alternative definitions of debt ratio. These definitions are: (a) long-term debt divided by the sum of long-term debt and book equity, and (b) total debt divided by book equity. The Fisher type tests (both ADF and PP) remain significant at levels higher than the 1% level. The CD test statistics range from 6.40 (Telecommunications Services) to 389.59 (Materials) in case (a), and from 12.10 (Materials) to 422.57 (Industrials) in case (b). The C_{AVE} statistics, however, range from 1.43 (Telecommunication Services) to 134.37 (Materials) in case (a), and from 6.85 (Utilities) to 139.67 (Industrials) in case (b). Finally, the CIPS* test statistics range from -1.15 (Telecommunication Services) to -1.89 (Consumer Staples and Utilities) in case (a), and from -1.48 (Materials) to -2.20 (Utilities) in case (b). The null hypothesis of the presence of a unit root can be rejected at the 1% level only in case (b) and for only the one sector, Utilities.

The Impact of the financial crisis on the stochastic properties of debt ratios

Do our results really invalidate the trade-off model? We argue that it is premature to make such a conclusion. Up to this point, we have assumed that throughout the entire sample period the stochastic process representation of debts ratios does not exhibit structural change. In the case, when this assumption fails, the tests can be misleading and biased toward the non-rejection of the unit root hypothesis. Thus, some caution should be exercised in interpreting our findings of non-stationarity under cross-sectional dependence because the global financial crisis and the resulting Great Recession are included in our sample. As shown by Dang et al. (2014), the speed of adjustment of corporate debt ratios has been significantly affected by the financial crisis. Using the dummy variable approach, Dang et al. (2014) find that the coefficient on the crisis dummy variable (which takes on value 1 if the year is between 2007 and 2009, and 0 otherwise), is negative and significant. Alternatively, using the sample-splitting approach, Dang et al. (2014) find that the estimate of the speed of adjustment for the

Table 6. Results of the $CIPS^{*}$ test for unit roots for both book value and market value debt ratios

Economic Sector	Book value debt ratio	Market value debt ratio
Materials	-1.521	-1.678
Consumer Discretionary	-1.558	-1.401
Consumer Staples	-1.686	-1.412
Health Care	-1.727	-1.449
Energy	-1.416	-1.510
Financials	-1.660	-1.268
Industrials	-1.395	-1.610
Information Technology	-1.904	-1.389
Telecommunications Services	-1.636	-1.454
Utilities	-2.208*	-1.849

Note: The truncated version of the CIPS, denoted as $CIPS^{*}$, is constructed to avoid the problem of an extreme statistic in cases when T is small, and is computed as $CIPS^{*}(N,T) = N^{-1}\sum_{i=1}^{N} t_i^{*}(N,T)$ where

$$t_i^{*}(N,T) = \begin{cases} t_i(N,T) & -K_1 < t_i(N,T) < K_2 \\ -K_1 & t_i(N,T) \le K_1 \\ K_2 & t_i(N,T) \ge K_2 \end{cases}$$

The parameters K_1 and K_2 are positive constants based on Monte Carlo simulations (K_1 = 6.19 and K_2 = 2.61). The limiting distribution of the $CIPS^{*}$ (N,T) statistic is non-standard, even in cases of a large N. The p-values are not available. The critical values are tabulated in Pesaran (2007) for various combinations of N and T. From Pesaran (2007) Table IIb (Case II: Intercept only), the critical values at the 1%, 5% and 10% levels are T = 15, N = 200: -2.16, -2.04, and -1.98; T = 15, N = 100: -2.19 and -2.07, and -2.00; T = 15, N = 50: -2.26, -2.11, and - 2.03, and denoted as. *, **, and ***, respectively.

period of the financial crisis is more than one half the corresponding estimate for the period that preceded the crisis. Empirical evidence also indicates that new lending declined dramatically during the financial crisis. Ivashina and Scharfstein (2010) document that new loans to large borrowers fell by 47% during the peak period of the financial crisis (fourth quarter of 2008) relative to the prior quarter and by 79% relative to the peak of the credit boom (second quarter of 2007). The decline is likely to reflect both demand and supply conditions. On the demand side, the drop in borrowing is the result of firms scaling back their expansion plans; on the supply side, banks more vulnerable to moral hazard problems tend to restrict the supply of loans. The combined effect of these two forces is likely to trigger a structural change that may potentially invalidate our findings.

In the presence of a known structural change, one approach would test for the unit root twice, before and after the break. In our case, splitting the full sample into two sub-samples, the first from 1997 to 2007, and the second from 2008 to 2010, and applying unit root tests to both subsamples, is virtually impossible, given the limited number of years in the second subsample. Accordingly, we only provide a robustness check by applying the same unit root methodology with the first subsample that ends in 2007.

The findings for the Fisher type tests do not modify the conclusions drawn for each sector using the original sample. The results of these tests, in both the ADF and the PP specifications, reject the unit root hypothesis for both the book value and the market value of the debt ratio series. Similarly, the findings of the CD and C_{AVE} remain robust to the sample reduction. That is, we find strong evidence of cross-section dependence in each sector for both measures of debt ratio. We do not report these findings, but they are available on request. However, some of the findings for the CIPS* are sensitive to the time period. We find more evidence of stationarity in the "pre-crisis" sample than in the

Table 6. Results of the $CIPS^*$ test for unit roots for both book value and market value debt ratios

Economic sector	Book value debt ratio	Market value debt ratio
Materials	-1.352	-2.235**
Consumer Discretionary	-1.232	-1.865
Consumer Staples	-1.587	-2.122**
Health Care	-2.068**	-1.618
Energy	-1.088	-1.487
Financials	-1.449	-1.323
Industrials	-1.402	-2.483*
Information Technology	-1.599	-1.493
Telecommunications Services	-0.904	-1.623
Utilities	-1.951	-2.227**

Note: The truncated version of the CIPS, denoted as $CIPS^*$, is constructed to avoid the problem of an extreme statistic in cases when T is small, and is computed as $CIPS^*(N,T) = N^{-1}\sum_{i=1}^{N} t_i^*(N,T)$ where

$$t_i^*(N,T) = \begin{cases} t_i(N,T) & -K_1 < t_i(N,T) < K_2 \\ -K_1 & t_i(N,T) \le K_1 \\ K_2 & t_i(N,T) \ge K_2 \end{cases}$$

The parameters K_1 and K_2 are positive constants based on Monte Carlo simulations ($K_1 = 6.19$ and $K_2 = 2.61$). The limiting distribution of the $CIPS^*$ (N,T) statistic is non-standard, even in cases of a large N. The p-values are not available. The critical values are tabulated in Pesaran (2007) for various combinations of N and T. From Pesaran (2007) Table IIb (Case II: Intercept only), the critical values at the 1%, 5% and 10% levels are T = 15, N = 200: -2.16, -2.04, and -1.98; T = 15, N = 100: -2.19 and -2.07, and -2.00; T = 15, N = 50: -2.26, -2.11, and - 2.03, and denoted as. *, **, and ***, respectively.

original sample. Furthermore, we find more evidence of stationarity for the market value debt ratio than the book value debt ratio. This outcome was to be expected because book values are largely unaffected by changes in stock prices. In Table 7, the CIPS* test rejects the null hypothesis of the unit root for the market value debt ratio in Materials, Consumer Staples, and Utilities at the 5% significance level, and Industrials at the 1% significance level. Similarly, the CIPS* test rejects the null hypothesis for the book value debt ratio in Health Care at the 5% significance level. Thus, at least one measure of the debt ratios in these sectors appears to exhibit a reversal in dynamics, from unit root to mean reversion. Because the results for the full sample indicate non-stationarity and the results of the reduced sample suggest stationarity for these five sectors, we conclude that in these sectors the financial crisis has destabilized debt ratios, switching the dynamics of the debt ratios from a mean reversion behavior to a random-walk dynamics. We interpret these results as indicative that for these sectors the recent global events may have triggered a structural break in the underlying data generation process. In the face of increased risk aversion by credit suppliers and widespread informational asymmetries, this outcome is not shocking. In such an environment, a pecking order may be generated, where retained earnings represent the least expensive source of financing. For the remaining sectors (Consumer Discretionary, Energy, Financials, Information Technology, and Telecommunication Services) instead, the financial crisis does not appear to have affected the unit root dynamics of debt ratios. The resilience of these sectors to the crisis may be indirect evidence that internal financing plays a non-trivial role in the determination of debt ratios. This lack of uniformity of our findings is not surprising and is consistent with the idea that the recent financial crisis has not had a homogeneous impact on the U.S. economy (Dang et al., 2014). Thus, overall, we find that the evidence on the debt ratios dynamics is mixed.

Conclusions

This paper is an empirical investigation of the dynamics of corporate capital structure. Do firms have target debt ratios? The literature on corporate capital structure suggests at least three possible mechanisms for explaining the determinants of debt ratios: the trade-off theory, the pecking order theory, and the market-timing theory. Existing empirical work has focused almost exclusively on the determinants of capital structure, and while they have produced substantial evidence on the relation between capital structure and its determinants, they have not been able to provide much evidence on the dynamics of debt ratios. This study brings new evidence to bear on this important issue. We approach the question from the viewpoint of the methodology of panel unit root tests and investigate whether debt ratios are mean reverting or alternatively exhibit a random-walk process. If the empirical findings provide evidence of stationarity, this is an indication that the dynamics of the debt ratios are mean reverting, and, consequently, firm financial behavior follows the trade-off theory. Otherwise, if the empirical results provide evidence of unit root dynamics, this signals that firm financial behavior evolves according to other theories of capital structure, such as the pecking order theory or the market-timing theory.

Employing a panel of 2,556 US public firms over the period 1997-2010, we investigate the stationarity properties of the book value and market value measures of debt ratios for ten economic sectors of the U.S. economy. We first employ Fisher type panel unit root tests and find evidence that is overwhelmingly favorable to a mean reversion, i.e., stationarity hypothesis, and, consequently, the trade-off theory. This finding is consistent with much of the literature. However, these first-generation tests rely on the assumption of cross-sectional independence. Our analysis provides evidence that this assumption is not supported by the data. Cross section dependence does matter and substantially affects the outcome of the tests. Thus, when we apply the second-generation panel unit root test developed by Pesaran (2007) that accounts for this dependence, the results challenge the notion that debt ratios are mean reverting. We view these findings as evidence that contradicts the trade-off theory, but is consistent with the pecking order and the market-timing theories. We perform a robustness check on our findings and consider whether the results of the panel unit root tests are sensitive to the selection of the sample period. We find that the recent macroeconomic events of the global financial crisis and the Great Recession play a crucial role in our understanding of the dynamics of debt ratios. We construct a pre-crisis sample that excludes the last three years at the end of our full sample. The results of this sample reduction generate more evidence of stationarity. Utilizing the market value debt ratio, four sectors (Materials, Consumer Staples, Utilities, and Industrials) exhibit stationary dynamics, while employing the book value debt ratio, one sector (Health Care) exhibits stationarity. We interpret these results as indicative that the recent global events may have produced in these sectors a structural change in the underlying data generation process (DGP).

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Appendix.

Description of the 10 sectors (Compustat economic sector codes in parenthesis).

Materials (1000) include all construction materials, chemicals, gases, and commodity firms.

Consumer Discretionary (2000) includes automobile manufacturers, homebuilders, hotels, casinos, retail, and electrical appliances firms.

Consumer Staples (3000) include food and drug retail and brewers.

Health Care (3500) includes health care and pharmaceutical firms.

Energy (4000) includes all types of oil and gas firms.

Financials (5000) include insurance, banking, and investment brokerage firms.

Industrials (6000) include conglomerates, construction, aerospace, and defense, heavy machinery, airlines, marine, trucking, railroads, and office services and supplies.

Information Technology (8000) includes information technology, software, electronics and semiconductor firms.

Telecommunication Services (8600) include network providers, broadband services, radio, television, and voice communication.

Utilities (9000) include electric, gas, water, and shipping firms.