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SECURITIZATION MARKETS AND CENTRAL BANKING: POLICY ANNOUNCEMENT EFFECTS

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ABSTRACT

European and US monetary authorities designed policies to revive the issuance of asset backed securities (ABS) after the financial crisis of late 2007 and keep consumer and business funding uninterrupted. I examine how policy announcements affected excess ABS spreads, that is, spreads of newly issued ABS over and above broader financial market spreads, in the context of event studies. Though policy announcements effectively reduced and stabilized excess ABS spreads by late 2009, the volume of new issuance remained significantly low.

JEL: E58, G1

KEYWORDS: Securitization, ABS Spreads, TALF, Central Banking

INTRODUCTION

The United States (US) subprime mortgage crisis in 2007 resulted in a sudden stop in worldwide demand and issuance of Asset Backed Securities (ABS). In response, central banks deemed it was necessary to keep the securitization markets alive so that money flows to financial institutions keep consumer and small business funding uninterrupted. To this end, the European Central Bank (ECB) and the US Federal Reserve (Fed) carried out programs to provide liquidity to issuers and investors of newly issued ABS, respectively. Campbell et al. (2011) find that the Fed's Term Asset Backed Securities Loan Facility (TALF) program announcements reduced excess ABS spreads in the secondary market. However, TALF targeted primary rather than secondary market spreads. Hence, in this study, I examine TALF announcement effects on primary market spreads following Campbell et al.'s event study methodology. In addition, I examine how ECB policy announcements of its credit provisions program affected ABS spreads in the European marketplace. To perform these tasks I examine a dataset of 107,692 ABS deals from Thomson One Banker of Thomson Reuters.

The strongest ECB policy announcement effect on excess ABS spreads is the expansion of financing eligibility to non-European Economic Area ABS issuers on May 25, 2007. ABS spreads (to swaps) were reduced by 86 basis points (bp) more than how broad financial market spreads (LIBOR-OIS spread) changed during the week of the announcement. The Fed's announcement on February 10, 2009 that TALF could possibly expand up to $1 trillion reduced excess ABS spreads by 13bp, while the last TALF ABS subscription on March 4, 2010 reduced them by 10bp. Results are robust to alternative specifications of ABS and broad market spreads. The findings suggest that these policy announcements successfully reduced and stabilized excess ABS spreads by late 2009. Nevertheless, new issuance of private-label ABS remained considerably low, and which, like other quantitative easing monetary policies, can be thought of as "pushing on a string".

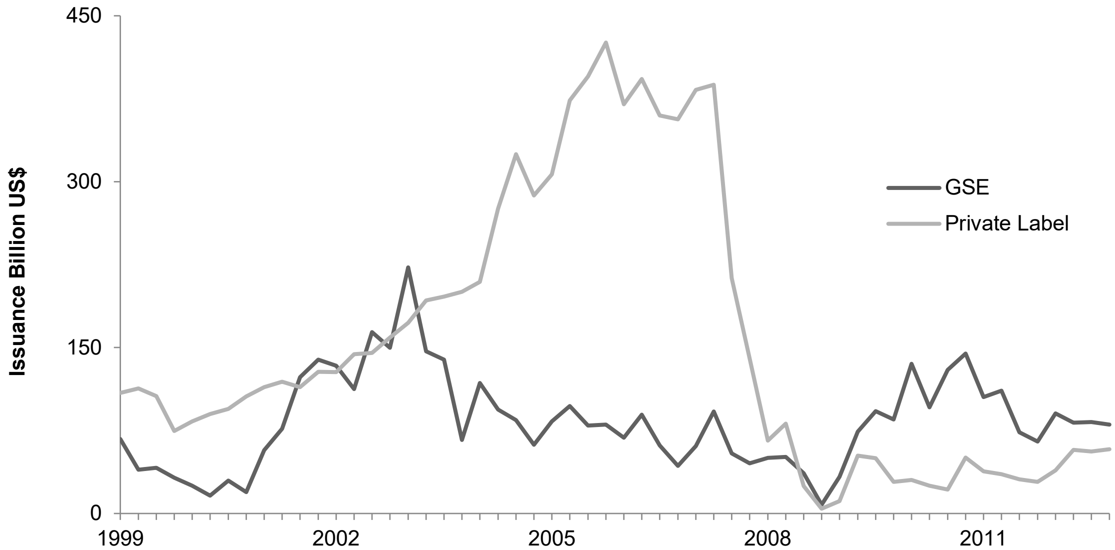
Background on Securitization

Securitization is the process of pooling diversified loans and bundling them in securities. The most common loan categories used in securitization are mortgage loans for Mortgage Backed Securities (MBS), and auto, credit card and student loans for Asset Backed Securities (ABS) (I use the term "ABS" with reference to both Mortgage and other Asset Backed Securities without loss of generality unless is otherwise necessary, in which case I make the distinction clear). Cash flows from loan repayments are used as collateral to repay the securities that are sold to investors such as insurance companies, hedge funds and pension funds. ABS are issued in different classes of credit-risk, called tranches. Senior tranches have lower credit risk and are considered safer investments than junior tranches. Senior tranches have repayment priority, pay a lower yield and receive higher investment-grade ratings than junior tranches. Typically, the most junior tranche stays with the originator so that it continues to monitor the original loan repayments. Cash proceeds from the sale of ABS to investors are then used to make new loans. Before August 2007, as much as 60% of private credit creation in the US was distributed onwards through the securitization market and was not held in the books of depository institutions (Dudley, 2009).

The origination of ABS is facilitated by a bankruptcy-remote robot firm, the Special Purpose Vehicle (SPV). The SPV is an off-Balance Sheet vehicle that allows the originator to remove the underlined assets from its balance sheet and lower its capital reserve requirements (Gorton and Souleles, 2005). The SPV finances the purchase of the pool of loans from the originator by issuing and selling ABS tranches to investors. Securitization originally developed in the US in the 1970s by Government Sponsored Enterprises (GSE) —such as Fannie Mae and Freddie Mac— for mortgage assets (MBS) and expanded in the mid-1980s to include non-mortgage assets (ABS) (Agarwal et al., 2010). Securitization later gained popularity among private financial institutions and industrial companies. Figure 1 shows that by year 2003 Private Label issuance from non-government agencies surpassed that of GSE. Since the Fall of Lehman Brothers in the third quarter of 2008, however, GSE issuance has averaged 70% of the total. At the same time that TALF launched, the Fed implemented a different program to buy up to $100 billion of GSE direct obligations and up to $500 billion of their MBS. The question this study seeks to answer is whether the TALF program was successful in reviving Private Label issuance. For this reason, I exclude GSE and investigate only newly issued Private Label ABS deals.

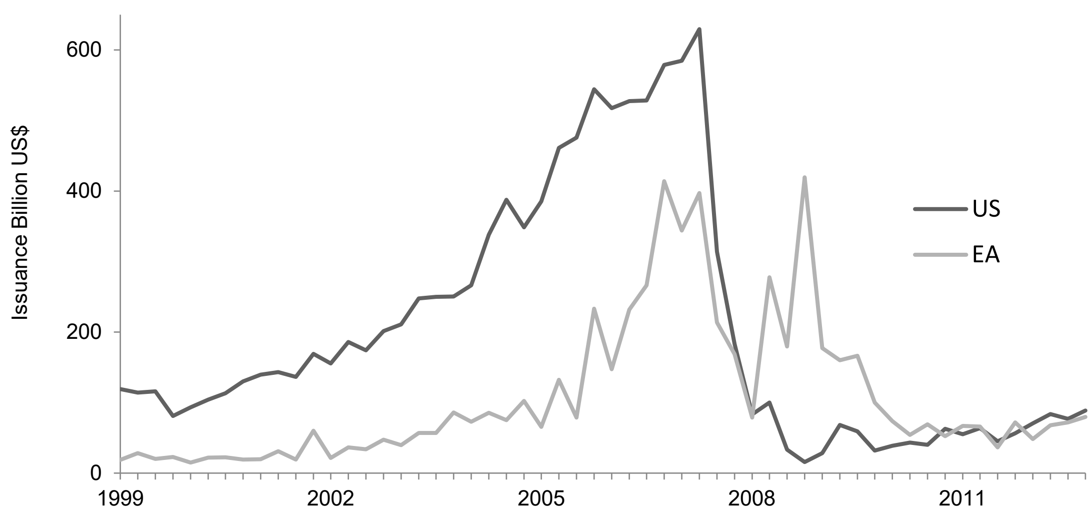
In Europe, securitization started to peak in 2002, when financial institutions began to collateralize European assets in addition to US assets. Figure 2 shows that the turmoil in the US housing market allowed the volume of issuance in the euro area marketplace to soar above US marketplace levels between the second quarter of 2008 and the third quarter of 2010. By 2012, however, issuance in the US marketplace was again higher than in the euro area marketplace, albeit significantly lower volumes in both markets relative to those years prior to 2007. Table 1 presents issuing patterns of Private Label ABS between 1999 —the year the euro was introduced— and 2012. Issuance by nations in the euro area, Great Britain and Switzerland accounted for more than 5.4 trillion dollars, or 35% of the total. Interestingly, the domestic and foreign currency denomination patterns of ABS deals vary significantly between countries but I leave that investigation for future study. The remainder of the paper is organized as follows. In the next section I review the relevant literature. Afterwards, I discuss the data and methodology, and present the timeline of the Fed and ECB program announcements. I then present the empirical results and finally conclude with suggestions for future research.

Figure 1: ABS Issuance in the US: Government Sponsored Enterprises (GSE) vs. Private Label Issuers



This figure shows issuance of Asset and Mortgage Backed Securities (ABS) in the US marketplace between 1999 and 2012. Government Sponsored Enterprises (GSE) refers to issuance by Fannie Mae, Freddie Mac, Government National Mortgage and SLM Corporation, all with standard industry classification code (SIC) 6111. Private Label refers to non-government sponsored institutions such as Bank of America and Lehman Brothers Holdings Inc. The volume of issuance is reported in billions of US dollars. Data source: Thomson One Banker.

Figure 2: Private Label ABS Issuance: US and Euro Area Marketplace



This figure shows issuance of Asset and Mortgage Backed Securities in the US and Euro Area (EA) marketplace between 1999 and 2012. The volume of issuance is reported in billions of US dollars. Data source: Thomson One Banker.

Table 1: Private Label Issuance by Originator Nation

	Total Issuance (1999-2012)		Currency Denomination (%)						
	(USD million)	(% of total)	Domestic	USD	EUR	GBP	JPY	CHF	Other
United States	8,581,438	55.72%	97.57%	…	1.33%	0.89%	0.04%	0.05%	0.12%
Euro Area countries	2,969,645	19.28%	80.97%	15.74%	…	2.57%	0.20%	0.03%	0.49%
United Kingdom	1,989,347	12.92%	45.79%	24.92%	28.25%	…	0.11%	0.09%	0.84%
Japan	623,004	4.05%	70.37%	27.73%	1.06%	0.79%	…	0%	0.05%
Switzerland	504,017	3.27%	0.32%	95.26%	3.56%	0.21%	0.08%	…	0.89%
Australia	346,472	2.25%	66.14%	21.88%	10.01%	1.83%	0.07%	0.00%	0.07%
South Korea	124,107	0.81%	73.12%	23.72%	2.50%	0.00%	0.52%	0.00%	0.14%
Canada	120,468	0.78%	70.52%	26.37%	2.44%	0.19%	0.24%	0.00%	0.24%
All other countries	86,704	0.92%	46.5%	30.92%	19.58%	1.29%	0.29%	0.00%	1.42%
Euro Area Countries Breakdown									
Spain	659,746	22.22%	86.44%	9.04%	…	4.15%	0.08%	0.07%	0.22%
Germany	518,656	17.47%	48.06%	49.65%	…	1.91%	0.00%	0.00%	0.38%
Italy	467,239	15.73%	93.52%	6.11%	…	0.25%	0.12%	0.00%	0.00%
Netherlands	454,383	15.30%	88.84%	7.07%	…	2.92%	0.00%	0.00%	1.17%
Ireland	362,367	12.20%	81.65%	10.49%	…	6.07%	0.75%	0.10%	0.94%
France	205,171	6.91%	75.14%	21.65%	…	0.95%	1.02%	0.09%	1.15%
Belgium	184,826	6.22%	98.01%	1.91%	…	0.00%	0.00%	0.00%	0.00%
All other EA	117,256	3.95%	96.34%	2.97%	…	0.69%	0.00%	0.00%	0.00%

This table shows issuance of Asset and Mortgage Backed Securities by private financial and industrial entities (Private Label ABS) for the period 1999-2012. For each country, the volume of total issuance is reported in millions of US dollars and is then broken down by ABS currency denomination. The currencies considered are the US dollar, European euro, British Pound, Japanese Yen, and Swiss Franc. All other currency choices are grouped together. Data source: Thomson One Banker

LITERATURE REVIEW

This paper relates to other policy event studies concerned with the global financial recession and that focus on yield spreads. A yield spread is the difference between the yield to maturity of a risky bond and a benchmark riskless bond of the same maturity. Spread data have been associated with business cycle fluctuations in the post-World War II era, as they tend to widen shortly before the onset of recessions and narrow again before recoveries (Gilchrist et al., 2009, Guha and Hiris, 2002). Rising spreads may reflect a decline in economic fundamentals due to a reduction in the expected present value of corporate cash flows prior to a downturn (Philippon, 2009), or the deterioration of corporate and financial intermediaries' balance sheets that disrupts the supply of credit (Bernanke et al., 1999, Adrian and Shin, 2010). Hence, the information in yield spreads contains both credit and liquidity risk premia, and may be indicative of how financial prices affect the real economy (Guidolin and Tam, 2013).

Several papers investigate the effectiveness of policy programs during the global financial crisis on market spreads (he crisis-related programs by the Federal Reserve can be divided into three groups. The first group is about the role of the central bank as a lender of last resort to provide short-term liquidity to financial institutions. These programs were the discount window, the Term Auction Facility (TAF), the Primary Dealer Credit Facility (PDCF), the Term Securities Lending Facility (TSLF) and the bilateral currency swap agreements with 14 foreign central banks. The second group of programs provided liquidity directly to borrowers and investors in key credit markets. These programs were the Commercial Paper Funding Facility (CPFF), the Asset Backed Commercial Paper Money Market Mutual Fund Liquidity Facility (AMLF), the Money Market Investor Funding Facility (MMIFF) and the Term Asset Backed Securities Loan Facility (TALF). The third group of programs is about expanding open market operations to support credit markets and put downward pressure on long-term interest rates. Such programs included the purchase of Government Sponsored Enterprise debt and Mortgage Backed Securities). Adrian et al. (2010) and Dwyer and Tkac (2009) find that the Commercial Paper Funding Facility (CPFF) was generally effective in averting a run on money market funds and easing liquidity into the commercial paper funding markets. Baba (2009) finds that the Fed's currency swap lines met the demand needs for US dollar from money market mutual funds.

Various studies find that the Term Auction Facility (TAF) successfully reduced Libor-Fed spreads (Ceccheti, 2009, Christensen et al. 2009, Frank and Hesse, 2009, Sarkar, 2009, and Wu, 2001) though Taylor and Williams (2009) do not. Hancock and Passmore (2011) find that the Fed's MBS purchase program reduced risk premiums in mortgage rates. Ashcraft et al. (2011, 2012) find that the TALF program reduced spreads of legacy (previously issued) CMBS that were accepted as TALF collateral, but only by a small amount. Campbell et al. (2011) find that the TALF program reduced secondary market spreads without subsidizing individual securities, and at a very low loss risk to the US government. Contrary to these studies, I evaluate the effects of these programs on primary market spreads (at issuance) rather than on secondary market spreads, as the explicit goal of the Fed's TALF and the ECB's credit provisions programs were to boost new ABS issuance that came to a halt. Secondary market spreads are not based on actual transactions. Instead, they depend on dealer quotes and pricing matrices. They can be affected by the overall stance of the economy and not just by ABS-related factors, which makes identification of policy effects solely on the ABS market difficult. As Vink and Fabozzi (2009) note, reliable secondary market spread data is typically challenging to obtain, while issuance spreads are a more accurate measure for the actual cost of debt and risk premium demanded by investors. Primary ABS spreads have previously been studied by Adelino (2009). He finds that spreads (other than AAA-rated) can predict future downgrades and defaults, even after controlling for the initial credit rating. Vink and Fabozzi (2009) find that non-US ABS spread determinants at issuance can be explained by credit ratings and bond market conditions. Obviously, the Fed and ECB policy announcements are expected to have a positive impact on the targeted primary market ABS spreads. This study, however, examines the impact of these programs on "excess" ABS spreads, that is, how much more ABS spreads have been affected relative to broad market spreads during the weeks of policy announcements. The methodology draws on Campbell et al. (2011).

DATA AND METHODOLOGY

Event Study Methodology

Event studies have been widely used to assess short-run response to policy announcements (See Campbell et al. (1997) and Kothari and Warner (2007) for the econometrics of event studies used in Finance). Their main strengths are simplicity and parsimony. An application to short-term changes of excess ABS spreads avoids the need to model the time-varying properties of its level such as trends, structural breaks, nonlinearities and non-stationarity. In addition, focusing on short time intervals allows one not to specify an underlying spread model. Although the basis of policy evaluation is narrow in event studies, they can still be suggestive of policies' long-term effectiveness. Positive immediate market reactions may be self-fulfilling and lay the ground for sustained policy success (Ait-Sahalia et al., 2012).

Nevertheless, event studies have certain limitations. Most importantly, they cannot control for all the possible factors that market participants may be reacting to and that occur simultaneously with the policy under study. An event study will assign any market response solely to the specific program's announcement. The caveat applies to this event study as well, as central banks carried out a series of programs simultaneously in response to the crisis. In addition, the timing of policy announcements and the market reaction to them is state-contingent. In the current context, the collapse of Lehman Brothers must have certainly made the perceivable appropriateness of policy programs and market reaction to them stronger than if such event had not occurred. I try to ameliorate these problems by considering alternative measures of financial market distress that capture market participants' responses to risk and liquidity, assessing in this way the robustness of the results. I also keep the event window short to a weekly frequency. The intuition of an event study on ABS spreads is to gauge whether program announcements reduce ABS spreads more than they reduce broader market spreads. If that is the case, then this study provides suggestive evidence that the programs had a positive liquidity effect on the ABS market during the week of an announcement.

The identification assumption is that the announcements come as a surprise to market participants, and in the absence of TALF announcements ABS spreads during each event window would be unchanged relative to broad market indices (Campbell et al., 2011). I employ event study techniques to assess the impact of TALF and ECB program announcements on newly issued ABS spreads in the US and euro area marketplaces, respectively. The procedure follows Campbell et al. (2011) and involves the following steps: (i) construct ABS spreads, (ii) proxy general (broad) market spreads, (iii) obtain ABS excess spreads as the difference of ABS and general market spreads and, (iv) regress weekly excess ABS spreads changes on program announcement dummies to evaluate how each announcement affected ABS spread weekly changes over and above broad market spread weekly changes.

Data Description

The ABS dataset used in this paper is obtained from Thomson One Banker Deal Analytics. It consists of micro-level characteristics of ABS deals based on actual transactions on the date of issuance. Deal data fields contain security identifiers for the Originator entity and Special Purpose Vehicle (SPV), asset type description (e.g. mortgage, credit-card, auto-loans), collateral nation (origin of asset type), principal amount, coupon rate type (fixed, floating), maturity date, average life, the marketplace in which the deal is offered for sale, and credit ratings from Standard & Poor's, Fitch and Moody's, where available. The original dataset comprises of 107,692 ABS deals in the years between 1999 and 2012. To my knowledge, this newly constructed dataset has not been previously used in ABS policy evaluation studies. Data on the overall stance of the economy, such as interest and credit default swap rates, is obtained from Datastream.

I drop US Government Sponsored Enterprises (GSE) deals from the original dataset and focus on the effects of the TALF program on the private sector (Government Sponsored Enterprises (GSE) refer to the US Federal and federally sponsored entities identified by the Standard Industry Classification (SIC) code 6111. They include Fannie Mae, Freddie Mac, Government National Mortgage, Farmer Mac and PA Higher Education Assistance). This amounts to ABS issuance originating from private financial and industrial institutions, or Private Label. I then drop deals with low credit ratings that would not be accepted as eligible collateral by either the Fed's TALF or the ECB's credit provisions program. Unfortunately, I cannot explore differences between the eligible and non-eligible ABS deals because non-AAA rated ABS tranches almost disappear during the post-Lehman panic, due to investor risk aversion. The sporadic and non-eligible deals vary significantly in terms of size, rating, maturity, collateral asset type and coupon rates. Thus, it would be unwise to consider them jointly as a control group and compare against the eligible (treated) group. I also drop floating-rate securities that are indexed over the LIBOR rate. I examine only fixed coupon-rate securities that have a pre-specified rate of return at the date of issuance. I do not consider Collateralized Debt Obligations (CDO) ABS whose rate of return depends on managerial performance. Finally, I winsorize the remaining deals using the coupon variable at the 1% level to eliminate observations that are clear mistakes in data entries, following Adelino, (2009). This leaves 12,839 ABS deals in the US marketplace and 776 in the euro area marketplace. The limited number of ABS deals in the euro area marketplace during the crisis does not permit a time series analysis of announcement effects on any single asset class. I assume that the ABS deal selection criteria outlined previously can reasonably minimize the differences across asset classes and enable the analysis of ECB policy announcement effects on the group of all eligible, newly issued, investment-grade, fixed coupon-rate ABS asset classes jointly. In the US marketplace, only auto ABS deals have frequent issues in the post-Lehman era. Hence, I evaluate the Fed's TALF program announcement effects solely for this asset class.

ABS Yield Spreads

I construct the ABS (yield) spread, s_t^i, for an ABS deal i issued at date t, as the difference of its fixed coupon rate and the swap interest rate of the closest maturity at the day of issuance. AAA-rated ABS have an average life of two years. Hence, if the ABS deal is offered for sale in the US marketplace, I subtract the two-year US interest rate swap from the coupon rate, and if the ABS deal is offered for sale in the euro area marketplace, I subtract the two-year euro interest rate swap (Datastream mnemonics for the two-year swap rates are ICUSD2Y for the US dollar and TREUR2Y for the Euro). The choice of ABS spreads to swaps is based upon market practices. For example, the European Banking Federation (EBF) considers the swap index to be the derivative market's reference rate for the euro. In addition, Campbell et al. (2011) obtain indicative dealer quotes on secondary market ABS spreads from J.P. Morgan trading desk that are also reported as spreads to swaps. ABS spreads could alternatively be specified over benchmark government bond yields or AAA-rated corporate bond yields. The correlation coefficient between the weekly changes of ABS spreads to swaps and these alternatives ranges between 0.952 and 0.998 in the period considered for the event studies. This indicates that results reported for ABS spreads to swaps would be qualitatively the same had these alternative spreads been used instead.

Broad Market Spreads

The primary proxy variable to assess broader market conditions is the LIBOR-Overnight Index Swap (LIBOR-OIS) spread, s_t^M, which is the spread between the three-month unsecured interbank borrowing rate (LIBOR) and a proxy for the risk-free rate, the three-month overnight index swap rate (OIS). The LIBOR rate increases as market conditions deteriorate for two reasons. The first is counterparty risk, as the borrowing bank's probability of default increases. The second is liquidity risk, as the fear of declining asset prices increases bank capital requirements and leads banks to precautionary hoard cash. On the other hand, the fixed OIS rate is considered risk-free for the following reasons. First, it is set by central banks and as such, it is relatively stable (The Overnight Index Swap (OIS) rate is the federal funds target rate for the US and the Euronia rate for the euro area. The respective Datastream mnemonics for these two series are OIEUR3M and USFDTRG). Second, in an OIS transaction the loan amount is notional, so no money is changing hands. Instead, the two parties agree to exchange only the accrued interest differential that results from a floating versus a fixed interest rate loan, at loan maturity. In case of default, the loss is limited to the accrued interest spread. Thus, OIS is considerably less risky than an actual LIBOR loan. Still, there is interest rate risk, reflecting the uncertain expectations of how interest rates will move in the future. Subtracting OIS from LIBOR will remove interest rate expectations inherent in all interest rates, so liquidity and credit risk are the two remaining components of the LIB-OIS spread.

LIBOR-OIS spread has been widely used as a measure of counterparty risk and/or liquidity in the banking system (Gorton, 2010, Gorton and Metrick, 2012, McAndrews et al., 2008, Taylor, 2009, Taylor and Williams, 2009, Wu, 2011). A natural alternative to LIBOR-OIS for general market spreads would be the TED spread, defined as the difference between the three-month LIBOR and the three-month Treasury bill rate. As Wu (2011) notes however, the T-bill rate is not a good proxy for the risk-free rate in periods of financial turbulence, as "flight to quality" increases the demand for Treasury bills more than in normal conditions. Campbell et al. (2011) alternatively proxy broader market conditions with indicative quotes on the 5-year CDX Series 9 index of investment grade corporate credit default swaps (CDS) from Markit. I consider all three variants of general market spreads to assess the robustness of the results for the TALF program evaluation. For the ECB program evaluation, I consider a fourth alternative, the three-month LIBOR-Repo spread. Financial institutions in Europe widely use repurchasing agreements (repo market) with the ECB to borrow against ABS collateral. Gorton and Metrick (2012) consider the financial crisis of 2007-09 as a bank run on the repo market.

Excess ABS Spreads

To get the excess spreads of a particular ABS deal i, I subtract the general market spread (e.g. LIBOR-OIS spread), s_t^M, from the ABS spread, s_t^i, at the day of issuance. I then obtain weekly average excess spreads, W_t, for all n AAA-rated fixed-coupon ABS deals as shown in Equation 1:

$$W_t = \frac{1}{n}\sum_{i=1}^{n}\left(\mathrm{s}_\mathrm{t}^\mathrm{i} - \mathrm{s}_\mathrm{t}^\mathrm{M}\right) \tag{1}$$

In weeks where there is no ABS issuance, I use linear interpolation to fill in the missing values of excess ABS spreads. Lastly, weekly excess ABS spreads changes, ΔW_t, form the dependent variable in a regression on program announcement dummies. Dummies take the value 1 in the week of the announcement and are zero otherwise. The interpretation of the dummy coefficients then is, how much more ABS spreads to swaps changed relative to broad market spreads from the previous week. A negative coefficient of 10 indicates that the TALF program announcement (or the ECB announcement in the case of the euro area marketplace study) reduced weekly ABS spreads by 10 basis points more than how weekly broad market spreads changed, relative to the previous week. The regression model is shown in Equation 2:

$$\Delta W_t = b_0 + b_1 An_1 + b_2 An_2 + \cdots + b_n An_n + \varepsilon_t \tag{2}$$

Securitization Policies

The Fed's Term Asset Backed Securities Loan Facility (TALF) Program

In the United States, housing prices started to fall in mid-2006, but the severe market shock occured when Lehman Brothers –a colossus investment bank– closed its subprime lender BNC Mortgage LLC on August 22, 2007 and later filed for Chapter 11 bankruptcy on September 15, 2008. Noting that new issuance of MBS and ABS sharply declined, The Federal Reserve Board noted the sharp decline in ABS issuance and responded by announcing the Term Asset Backed Securities Loan Facility (TALF) on November 25, 2008. To increase investor demand for newly issued ABS, the Federal Reserve Bank of New York (FRB-NY) would lend up to \$200 billion to holders/investors of AAA-rated ABS backed by newly and recently originated consumer and small business loans. The loans were on average 90% of the ABS collateral value. The remaining 10% of the ABS value would have to come out of investors' own funds, called the loan "haircut" (Adrian and Shin (2010), Ashcraft et al. (2011), Brunnermeier and Pedersen (2009), and Gorton and Metrick (2009) study the role of increasing haircuts in the Panic of 2007-2008). Loans were non-recourse, which means that investors could put an ABS back to the Fed if its value declined. Thus, the maximum amount investors could potentially lose was the loan haircut.

On December 19, 2008, the Federal Reserve released detailed operational aspects of the TALF loans: eligible collateral included student, auto, credit card, and loans guaranteed by the Small Business Administration (SBA); loan maturity extended from one to three years, and loans would be provided to all borrowers with eligible collateral rather than through an auction. On February 6, 2009 the Fed released additional terms and conditions of the TALF, including loan rates and collateral haircuts. On February 10, 2009, the Board of Governors announced that the size of the TALF could increase up to \$1 trillion. In addition, eligible collateral expansion was under review to include newly issued AAA-rated commercial mortgage-backed securities (CMBS), private-label residential MBS, and other ABS. The program launched on March 3, 2009. The first funds from the TALF disbursed on March 25, 2009 with student, auto, credit card, and SBA guaranteed loans being the eligible ABS asset classes. The program would make monthly funding subscriptions until December 2009. On March 19, 2009, the set of eligible collateral for the April 2009 subscription expanded to include ABS backed by mortgage servicing advances, leases relating to business equipment, leases of vehicle fleets and floor-plan loans.

On March 23, 2009, the Federal Reserve announced that it considered including certain legacy (previously issued) securities in the list of eligible asset types for TALF loans. On April 21, 2009, the FRB announced two new interest rates for fixed-rate ABS with weighted average lives to maturity (WALM) of less than two years. On May 1, 2009, the FRB announced that commercial MBS (CMBS) and insurance premium loans would become eligible collateral under the TALF, and loan maturities would increase from three to five years starting in June. On May 16, 2009, the first new-issue CMBS subscription was carried out. On May 19, 2009, the Fed announced that starting in July, certain high-quality MBS issued before January 1, 2009 (legacy CMBS) would also become eligible collateral. These CMBS would be reviewed by the FRBNY, had to have at least two triple-A ratings from credit rating agencies (CRAs), and be the most senior in payment priority to all other tranches of an ABS deal.

On May 26, 2009, Standard and Poor's (S&P) CRA announced its intention to change its rating procedures in a way that would significantly reduce the amount of Legacy MBS eligible for TALF loans. The initial subscription date for TALF loans collateralized by newly issued CMBS was set for June 16, 2009. In June 25, 2009, the Federal Reserve announced extensions and modifications to a number of its liquidity programs, but the TALF expiration date remained unchanged to December 31, 2009. On June 26, 2009, S&P proceeded with their proposed changes, putting many AAA-CMBS in downgrade watch. The first legacy CMBS subscription took place on July 16, 2009. On August 17, 2009, the Federal Reserve and Treasury approved an extension of the TALF program through March 31, 2010 for ABS and through June 30, 2010 for newly issued CMBS. On October 5, 2009, the Federal Reserve proposed rules to promote competition among Nationally Recognized Statistical Rating Organizations (NRSRO) that provided credit ratings for ABS. It also announced that starting with the November 3 subscription, the FRB-NY would conduct a formal risk assessment of all proposed collateral, in addition to requiring AAA-ratings from two NRSRO. On November 17, 2009, the first newly issued CMBS deal took place. On December 4, 2009, the FRB announced the adoption of the October 5 proposal on the NRSRO eligibility for credit ratings. The final ABS, legacy CMBS and new-issue CMBS subscription dates were on March 4, March 19, and June 18, 2010, respectively.

When the TALF program closed on June 30, 2010, only $70 billion out of the $200 billion that the FRB authorized were extended in TALF loans, and only $43 billion were outstanding. Despite loan maturity ranging between three to five years, many loans were repaid early due to the interest rate structure of the TALF loans; they were lower than the market rates in the midst of the financial crisis in 2008 but higher than in normal times, giving investors an incentive to repay the loans once financial markets normalized.

<u>The European Central Bank's Eligible Collateral Program for Credit Provisions</u>

National central banks and financial institutions in Europe can borrow from the Eurosystem's credit operations only on a repurchase agreement (repo) collateralized basis (Repurchase agreements (repos) are collateralized lending transactions. One party agrees to sell securities to the other against a transfer of funds. At the same time, the parties agree to repurchase the same or equivalent securities at a specific price in the future. Consider this hypothetical example: the market value of the assets backing an ABS is $100 billion, but the ABS is issued at $80 billion. The eligible issuer can borrow $80 billion against the ABS from the Eurosystem and agree to repurchase it for $88 billion. In this example, the ABS is over-collateralized by a 20% haircut/margin, and the repo haircut/rate is 10%). Assets accepted as collateral must fulfill certain criteria that the European Central Bank (ECB) publicizes and frequently updates. It is worth noting that unlike the US Fed's TALF program which allotted money to investors of ABS, the ECB program allotted money to the issuers of ABS. To my knowledge, this is the first event study to examine how ECB's changes in eligibility criteria for Asset Backed Securities (ABS) affected the issuance of newly issued ABS in the euro area marketplace.

On January 13, 2006, Asset Backed Securities (ABS) were accepted for the first time as collateral if they were issued by an entity established in the European Economic Area (EEA), had at least a single-A credit rating or higher, and were denominated in euros (The European Economic Area consists of the 27 countries of the European Union, plus Iceland, Liechtenstein and Norway). ABS securities originated by entities based in the United States, Canada, Japan, or Switzerland were not eligible until May 25, 2007.

The same credit rating and euro-currency denomination criteria applied to these deals once they became eligible. On October 15, 2008, the list of eligible assets temporarily expanded to include ABS issued in three currencies other than the euro, but only if they were issued in the euro area marketplace. Specifically, these currencies were the US dollar, the British pound and the Japanese yen. The temporary measures were to remain in effect until the end of 2009, but the operative date would be determined later. These instruments were subject to a uniform haircut add-on of 8%. On the same date, the Eurosystem started to offer US dollar liquidity through foreign exchange swaps (Goldberg et al (2011) review the literature on the Temporary Reciprocal Currency Arrangements that the Federal Reserve established with fourteen foreign central banks (dollar swap lines) between 12/2007 and 02/2010). On November 12, 2008, the ECB announced that non-euro denominated ABS would become eligible collateral as of November 14, 2008. As outlined, these securities ought to be issued by an issuer established in the EEA, be held or settled in the euro area, and fulfill all other eligibility criteria as set by previous decisions.

On January 20, 2009, the Eurosystem raised the bar on credit ratings requirements, requiring a triple-A (AAA/Aaa) rating up from a previous single-A rating from an accepted external credit assessment institution (ECAI) at issuance, for all ABS issued as of March 1, 2009. Previously issued (legacy) ABS would have to retain the single-A rating requirement over the lifetime of the ABS. On May 7, 2009, the Governing Council of the ECB decided to prolong the end of the temporary expansion of the list of eligible assets, from year-end 2009 to year-end 2010.

Table 2: Announcements Considered in the ECB and TALF Program Event Studies

Date	Program	Announcement Description
01/13/2006	ECB	ABS eligible for Eurosystem credit operations must be denominated in euros, issued by entities established in the European Economic Area (EEA) and have at least one A-rating from a credit rating agency (CRA). Currently excludes the United States, Canada, Japan and Switzerland.
05/25/2007	ECB	ECB expands eligibility to issuers of the four non-EEA G10 countries: the United States, Canada, Japan and Switzerland. ABS must still be denominated in euros.
08/09/2007	ECB	ECB injects 94.8 billion euros in overnight credit into the interbank market, after the French Bank BNP Paribas froze redemptions for three investment funds, citing its inability to value structured products.
10/15/2008	ECB	ECB announces eligibility expansion to marketable debt instruments denominated in non-euro currencies, namely the US dollar, the British pound and the Japanese yen, and issued in the euro area. The Eurosysten starts offering US dollar liquidity through foreign exchange swaps.
11/12/2008	ECB	ECB announces that eligibility expansion to non-euro marketable debt instruments will start on November 14, 2008.
11/25/2008	TALF	Fed announces the TALF program
12/18/2008	TALF	TALF details announcement. Focus still on ABS, maturity of TALF loans extended from one to three years.
01/20/2009	ECB	ECB requires a triple-A rating from a CRA.
02/10/2009	TALF	TALF possible expansion to $1 trillion. CMBS mentioned as a possible collateral type. Under the Treasury's Financial Stability Plan, the Treasury would use $100 billion to leverage up to $1 trillion in lending (up from $20 billion and $200 billion respectively).
03/03/2009	TALF	TALF launches. First consumer ABS TALF subscription announcement
08/17/2009	TALF	TALF extension until March 31, 2010 for Asset Backed Securities
11/20/2009	ECB	ECB requires two triple-A ratings from CRAs
03/04/2010	TALF	Last commercial ABS subscription
03/19/2010	ECB	ECB announces that as of January 1, 2011, non-euro denominated ABS will no longer be eligible collateral
06/30/2010	TALF	TALF closes

This table shows the timeline and description of ABS policy announcements considered in the event studies. The ECB program refers to the European Central Bank's Eligible Collateral program for credit operations with the Eurosystem. TALF refers to the Fed's Term Asset Backed Securities Loan Facility (TALF) program.

On November 20, 2009, the Eurosystem required at least two triple-A ratings from an accepted ECAI for all ABS issued as of March 1, 2010. On April 10, 2010, the ECB confirmed that the US dollar, pound sterling and Japanese yen denominated marketable debt instruments issued in the euro area would no longer be accepted collateral as of January 1, 2011. On April 23, 2010, the Eurosystem launched preparatory work on the establishment of loan-level information requirements for ABS deals in its collateral framework. The goal was to increase transparency and restore confidence in the market. Evidently, the ECB's eligibility criteria not only depended on collateral asset classes, but to issuer demographics and currency denomination. Table 2 provides a short description of the announcements considered to have an effect on excess ABS spreads, for either the ECB or the TALF program event studies.

RESULTS AND DISCUSSION

In this section, I present the results of event studies on how the announcements of the Fed's Term Asset Backed Securities Loan Facility (TALF) and the ECB's eligible collateral program for credit operations with the Eurosystem affected ABS excess spreads in their jurisdictions.

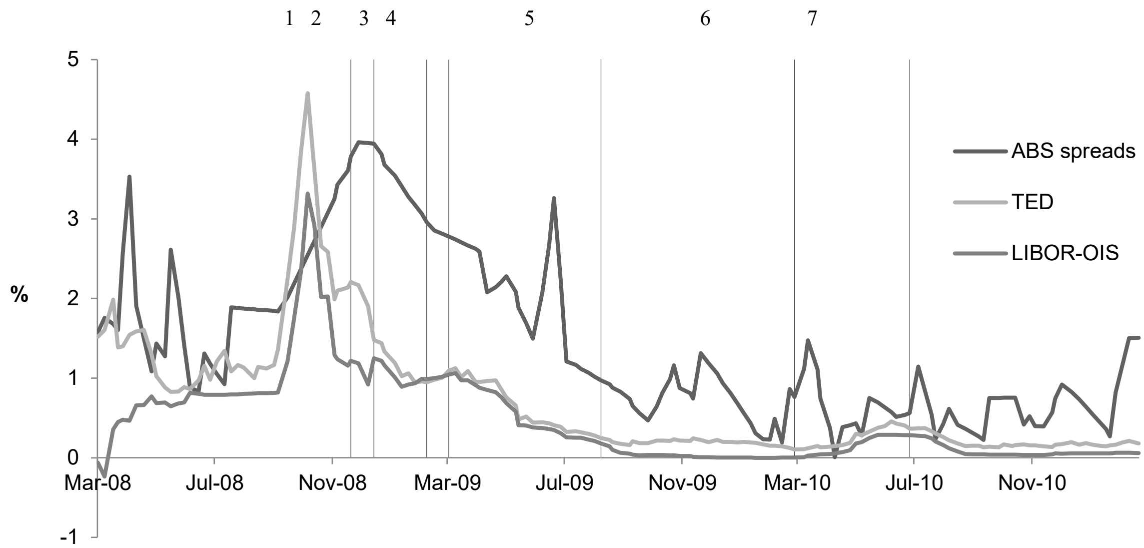
Term Asset Backed Securities Loan Facility (TALF) Program Evaluation

The sampling period for the TALF program evaluation is from March 4, 2008 to March 06, 2011. Excess spreads are constructed for AAA-rated, fixed-coupon auto ABS deals at the day of issuance. Then they are averaged at a weekly frequency as described in Equation 1. Right after Lehman Brother's collapse on September 15, 2008, there are no AAA-rated ABS deals in the Thomson One database until the announcement of the TALF program on November 25, 2008. Issuance has resumed since, but in low volumes and infrequent issues. Most of it comprises auto ABS. Private-Label issuance of Mortgage Backed Securities and of student loans practically disappeared in the period under study.

Credit-card issuance is present but with very infrequent deals, not allowing for a separate weekly horizon event study analysis for this specific asset class. For these reasons, the event study on the TALF program uses only the AAA-rated, fixed-coupon auto ABS asset class. Out of 156 weeks in the sample, excess spreads for 76 weeks are linearly interpolated since there are no auto ABS deals to match the criteria: Private-Label, fixed-coupon, AAA-rated, offered for sale in the US marketplace. Three alternative measures of general market spreads were considered: (1) three-month LIBOR-OIS spread, where OIS is the federal funds target rate; (2) three-month TED spread, the difference between LIBOR and the Treasury bill rate in the secondary market; and (3) the CDS weekly spread, the weekly change of the credit default swap index CMA CDX of investment grade securities in the Fall of 2007, in North America (Datastream mnemonic DCIG9S5). Figure 3 plots auto ABS spreads to swaps, the TED spread and the LIBOR-OIS spread for the period March 4, 2008 to March 6, 2011.

The LIBOR-OIS and TED spreads spiked in mid-October 2008. They began to fall thereafter as the $700 billion Treasury-run Troubled Assets Relief Program (TARP) restored liquidity in the broad financial markets capital through its capital injections. Note, however, that ABS spreads and counterparty risk continued to rise. This warranted the Fed to step in and design a program specifically for ABS, namely the TALF program. The first two announcements of the TALF program did not seem to affect ABS spreads, but spreads started to decline in early 2009. The picture is consistent with Figure 2 in Campbell et al. (2011) on secondary market spreads.

Figure 3: Auto ABS and General Market Spreads in the US Marketplace for the Period 03/2008 to 03/2011



This figure shows ABS excess spreads and general market spreads in the US marketplace for the period March 2008 to March 2011. ABS spreads are defined as the difference between newly issued, AAA-rated, fixed-coupon rates of auto ABS and two-year interest rate swaps. LIBOR-OIS is the spread between the 3-month unsecured US dollar interbank borrowing rate (LIBOR) and the federal funds target rate, a proxy for the risk-free rate; TED is the spread between the 3-month LIBOR and the 3-month Treasury bill in the secondary market. The vertical lines represent the seven program announcements considered in the TALF event study: (1) Initial TALF program announcement, (2) program details, (3) possible expansion to $1 trillion, (4) program launches its first ABS subscription, (5) program extension, (6) last ABS subscription, (7) program closure. ABS deal data is obtained from Thomson One Banker and general market spreads data from Datastream.

Table 3 presents the regression results of excess ABS spreads under the three alternative specifications of general market spreads. The initial announcement of the TALF program on November 25, 2008, did not reduce ABS spreads. In fact, ABS spreads continued to rise over all three alternative broad market spreads. Weekly changes of ABS spreads were 14 basis points (bp) higher than the weekly LIBOR-OIS spread changes, 9bp higher than the TED, and 7bp higher than the CDS, indicating that credit and liquidity risk in the US securitization market continued to rise. The second event date on December 19, 2008, specified which asset classes are eligible for TALF loans. ABS spreads rose 19bp over CDS spreads and by 32bp over TED spreads. Both are statistically significant at the 1% level. The spreads over LIBOR-OIS were 3bp lower, but not statistically significant.

The differences in these results arise because the TED spread was reduced in the period between the two announcements, while the LIBOR-OIS spread remained relatively constant, as we see in Figure 3. Since 2009, LIB-OIS and TED spreads co-move very closely. The next three announcements in year 2009 reduced excess ABS spreads under all three market spread specifications. The strongest announcement effect was the potential expansion of the program up to $1 trillion on February 10, with excess ABS spreads falling by 13bp over the LIBOR-OIS spread. Similar results were obtained under the other two measures. The first loan subscription on March 3 reduced excess ABS spreads by 3bp over LIBOR-OIS, 6bp over TED and 12bp over CDS spreads. The announcement on August 17, that the program would extend from year-end 2010 to March 2011, reduced spreads by 2-3bp but not with statistical significance. The last TALF ABS subscription in March 4, 2010 reduced excess ABS spreads by 10bp over LIBOR-OIS and TED spreads, or 3bp over CDS spreads. When the program closed in June 30, 2010, ABS spreads did not change significantly over general market spreads. Overall, the event study finds that the TALF program announcements effectively reduced ABS spreads of newly issued auto ABS relative to broad market spreads, by decreasing counterparty risk and increasing liquidity in the US securitization market.

Table 3: Changes in Excess Spreads of AAA-Rated Auto ABS Around TALF Policy Announcements

Announcement	**(1) LIBOR-OIS**	**(2) TED**	**(3) CDS**
11/25/2008, (first announcement)	0.14%***	0.09%***	0.07%**
	(4.93)	(3.06)	(2.18)
12/18/ 2008, (details)	-0.03%	0.32%***	0.19%***
	(-0.95)	(10.79)	(5.74)
02/10/2009, ($1 trillion expansion)	-0.13%***	-0.08%***	-0.11%***
	(-4.36)	(-2.81)	(-3.33)
03/03/2009, (launch)	-0.03%	-0.06%**	-0.12%***
	(-1.10)	(-2.03)	(-3.77)
08/17/2009, (extension)	-0.02%	-0.03%	-0.02%
	(-0.78)	(-1.12)	(-0.7)
03/04/2010, (last subscription)	-0.10%***	-0.10%***	-0.03%
	(-3.55)	(-3.28)	(-0.83)
06/30/2010, (closure)	0.03%	0.06%*	-0.05%
	(0.88)	(1.89)	(-1.56)
R^2	0.00	0.00	0.00
N	156	156	134

This table shows regression results for the Fed's Term Asset Backed Securities Loan Facility (TALF) event study. The estimated equation is $\Delta W_t = b_0 + b_1 An_1 + b_2 An_2 + \cdots + b_n An_n + \varepsilon_t$. *The sample is weekly and covers the period March 4, 2008 - March 6, 2011. The dependent variable is the weekly change of average excess auto Asset Backed Security (ABS) spreads. Excess ABS spreads are defined as the ABS spreads over general market spreads. ABS spreads are the coupon rates of newly issued, AAA-rated, fixed-rate auto ABS offered for sale in the US marketplace, minus the two-year interest rate swap. Three alternative general market spreads are considered. (1) LIB-OIS is the spread between the 3-month unsecured US dollar interbank borrowing rate (LIBOR) and the federal funds target rate, a proxy for the risk-free rate. (2) TED is the spread between the 3-month LIBOR and the 3-month Treasury bill in the secondary market. (3) CDS is the weekly change of the credit default swap index CMA CDX in North America of Investment grade securities in the Fall of 2007. ABS data is obtained from Thomson One Banker and general market spreads from Datastream. Each column corresponds to a different OLS regression that uses an alternative general market spread specification. Each row shows the coefficient estimate resulting from a regression of the relevant dependent variable on TALF announcement dummies. The coefficients capture how each announcement affected ABS spread weekly changes over and above broad market spread weekly changes. Results are reported in percentage points with t-statistics in parenthesis. The constant is omitted. ***,**, and * indicate statistical significance at the 1, 5 and 10 percent levels respectively.*

The findings are qualitatively similar to those of Campbell et al. (2011) who also find that excess ABS spreads start to decline after March 3, 2009. However, the magnitude of announcement effects is smaller for the actual, primary market ABS data used in this study relative to Campbell et al.'s on indicative, secondary market spreads. This study finds that newly issued ABS spreads on AAA-rated, fixed-coupon auto ABS deals were reduced from a high of 396bp when the TALF was announced to 76bp when the program closed. This translates into a reduction of excess ABS spreads over general market spreads (LIBOR-OIS) from 266bp at the start of the program to 28bp at the close of the program.

ECB Eligible Collateral for Eurosystem's Credit Operations Program Evaluation

The sampling period for the ECB program evaluation is from September 5, 2006 to September 6, 2010. Excess spreads are constructed for all AAA-rated, fixed-coupon ABS deals at the day of issuance. Then they are averaged at a weekly frequency as described in Equation 1. Out of 261 weeks in the sample, 43 weeks have no ABS deals that match the study's criteria: Private-Label, fixed coupon rate, AAA-rated, offered for sale in the euro area. For these 43 weeks, the values of excess spreads are linearly interpolated. Four alternative measures of general market spreads are considered: (1) LIBOR-OIS, the spread between the 3-month unsecured euro interbank borrowing rate (LIBOR) and the 3-month euro Overnight Interest Swap (OIS) rate, a proxy for the risk-free rate; (2) LIBOR-Repo, the spread between the 3-month LIBOR and the 3-month repo rate used in repurchase agreements; (3) TED, the spread between the 3-month LIBOR and the 3-month AAA-rated government bond redemption yield; (4) CDS, the weekly change of the credit default swap index series ITRXTSF5 for investment-grade, senior financial institutions in the euro area. Table 4 presents the regression results of excess ABS spreads under the four alternative specifications on eight ECB announcement dummy variables. Each dummy variable is equal to 1 on the week of the announcement, and is equal to 0 otherwise. The coefficient on each dummy is interpreted as the weekly change in ABS spreads relative to broad financial market spreads in the week of the announcement. ECB criteria initially supported ABS denominated in euros and issued by entities in the European Economic Area (EEA). The relevant announcement, on January 13, 2006, reduced ABS spreads by 6 basis points (bp) more than general market spreads.

Namely, the LIBOR-OIS and LIBOR-Repo spreads. However, this difference is not statistically different from zero. It is likely that market participants found the current eligibility criteria restrictive. The three-month TED spread and the iTRAXX CDS index were not available in early 2006 so no coefficient values are reported for specifications (3) and (4) in Table 4. The second announcement expanded eligibility to ABS issued by entities established in the US, Canada, Japan or Switzerland, in May 25, 2007. Here, excess ABS spreads are reduced by 86bp under the three-month LIBOR-OIS or LIBOR-repo spreads, by 89bp under the TED spread and by 87bp under the CDS spread. Out of all ECB policy announcements considered in this event study, this one had the strongest effect. The third announcement considered, on August 9, 2007, is not related to the ABS eligibility program but is added as control for a major event. ECB injected 94.8 billion euros in overnight credit in the interbank market after French bank BNP Paribas froze redemptions for three investment funds, citing its inability to value structured products (Brunnermeier, 2009). The LIBOR rate shot up at 440bp, but ABS spreads to swaps did not react much differently than our proxies for general market spreads. Excess spreads are 3, 2 and 5bp below the first three LIBOR related proxies, and 10bp above the CDS market proxy. None is statistically significant. This is a positive finding, as the underlying assumption in this event study is that in the absence of ABS related announcements, ABS spread changes should not be any different from broad market spread changes.

Table 4: Changes in excess ABS Spreads in the Euro Area Around ECB Policy Announcements

Announcement	(1)	(2)	(3)	(4)
	LIBOR-OIS	LIBOR-Repo	TED	CDS
01/13/2006, (€ currency)	-0.06%	-0.06%	-	-
	(-1.25)	(-1.28)		
05/25/2007, (US,CA,JP,CH issuers)	-0.86%***	-0.86%***	-0.89%***	-0.87%***
	(-17.36)	(-17.42)	(-13.10)	(-12.63)
08/09/2007,(€94.8 billion injection)	-0.03%	-0.02%	-0.05%	0.10%
	(-0.55)	(-0.32)	(-0.71)	(1.42)
10/15/2008, (non-€ announcement)	-0.22%***	-0.23%***	-0.29%***	-0.12%*
	(-4.46)	(-4.72)	(-4.28)	(-1.73)
11/15/2008, (non-€ start)	-0.22%***	-0.23%***	-0.29%***	-0.12%*
	(-4.46)	(-4.72)	(-4.28)	(-1.73)
01/20/2009, (AAA rating required)	-0.41%***	-0.41%***	-0.42%***	-0.46%***
	(-8.30)	(-8.30)	(-6.24)	(-6.74)
11/20/2009, (two AAA ratings required)	-0.34%***	-0.32%***	-0.40%***	-0.32%***
	(-6.87)	(-6.54)	(-5.85)	(-4.65)
03/19/2010, (only € as of 01/01/2011)	0.27%***	0.27%***	0.26%***	0.26%***
	(5.49)	(5.42)	(3.80)	(3.72)
R^2	0.01	0.01	0.01	0.01
N	261	261	184	178

This table shows regression results for the European Central Bank's (ECB) Eligible Collateral Program for Credit Provisions event study. The estimated equation is $\Delta W_t = b_0 + b_1 An_1 + b_2 An_2 + \cdots + b_n An_n + \varepsilon_t$. *The sample is weekly and covers the period September 6, 2005 to September 6, 2010. The dependent variable is the weekly change of average excess Asset Backed Security (ABS) spreads. Excess ABS spreads are defined as the ABS spreads over general market spreads. ABS spreads are the coupon rates of newly issued, AAA-rated, fixed-rate ABS offered for sale in the euro area marketplace, minus the two-year EONIA interest rate swap. Four alternative general market spreads are considered. (1) LIBOR-OIS, the spread between the 3-month unsecured euro interbank borrowing rate (LIBOR) and the 3-month euro Overnight Interest Swap (OIS) rate, a proxy for the risk-free rate. (2) LIBOR-Repo, the spread between the 3-month LIBOR and the 3-month repo rate used in repurchase agreements. (3) TED, the spread between the 3-month LIBOR and the 3-month AAA-rated government bond redemption yield; (4) CDS, the weekly change of the credit default swap index series ITRXTSF5 for investment-grade, senior financial institutions in the euro area. ABS data is obtained from Thomson One Banker and general market spreads from Datastream. Each column corresponds to a different OLS regression that uses an alternative general market spread specification. Each row shows the coefficient estimate resulting from a regression of the relevant dependent variable on ECB announcement dummies. The coefficients capture how each announcement affected ABS spread weekly changes over and above broad market spread weekly changes. Results are reported in percentage points with t-statistics in parenthesis. The constant is omitted. ***,**, and * indicate statistical significance at the 1, 5 and 10 percent levels respectively.*

On October 15, 2008, ECB announced that non-euro denominated ABS will become eligible asset classes, but the effective date was unknown. That same week the Eurosystem started to offer US dollar liquidity through foreign exchange swaps. On November 12, 2008, ECB announced the beginning of non-euro ABS eligibility. Both of these announcements had the same effects. Excess spreads over LIBOR related market spreads were reduced by 22 to 29bp, and by 11bp over the CDS spreads. The difference between the LIBOR related excess spreads findings and the ones over the CDS measure can be attributed to liquidity, as LIBOR market spreads capture both liquidity and credit risk, whereas the CDS measure only credit risk. The next two announcements raised the credit-rating eligibility requirements. The January 20, 2009 announcement required one triple-A rating up from a single-A rating, while the November 20, 2009 announcement required two triple-A ratings. These announcements reduced ABS excess spreads by 41 and 34bp over the LIBOR-OIS spread.

The negative sign on the coefficients is not surprising as the event study uses only the vintage of ABS that these announcements favored, the triple-A rated. Data on non-AAA rated securities are limited and spreads vary considerably across asset classes and credit ratings, making it impossible to assess and compare announcement effects with the "control" group of the program. The last announcement considered is on March 19, 2010, when ECB announced that the temporary expansion of non-euro eligible assets would stop as of January 2011. Excess ABS spreads rose by 27bp because of these negative news. Results are robust under alternative broad market conditions specifications. Coefficients on announcement dummy variables carry the expected sign and are statistically significant at the 1% for most ABS related announcements. This offers supporting evidence that the event study captures the direction and perhaps the magnitude of the short-run market response to ECB policy announcements, as reflected in the movement of newly issued excess ABS spreads.

CONCLUDING COMMENTS

This study follows the event study methodology of Campbell et al. (2011) to examine how the Fed's Term Asset Backed Securities Loan Facility (TALF) program announcements affected auto ABS spreads. Campbell et al. examine indicative, secondary market spreads whereas I examine primary market spreads of newly issued auto ABS based on actual transactions. I find that the Fed's TALF program announcements reduced and stabilized primary market spreads, consistent with the findings of Campbell et al., but my results are lower in magnitude. The most effective announcement was that the TALF program could expand to $1 trillion, on February 10, 2009. This reduced excess ABS spreads by 13 basis points. I also examine the ECB's announcement effects on a separate event study. In particular, the eligible collateral program for Eurosystem credit operations on the European ABS marketplace. I find that the announcement that lowered excess ABS spreads the most was on May 25, 2007, when the ECB extended eligibility to non-European Economic Area issuing institutions. ABS spreads fell by 86 basis points more than broad market spreads during the week of this announcement.

A closer look at the policy announcements from the Fed and the ECB reveals that the two central banking institutions might have been coordinating their policy responses. First, non-euro denominated securities in the euro area marketplace became eligible collateral at the same time when swap lines expanded between the two institutions and other European National banks, which aimed to provide European financial institutions with dollar liquidity. Second, the credit rating requirements for ECB eligibility were initially set to a single-A rating, but once the TALF program got underway specifying a triple-A rating requirement, the ECB followed suit and matched the criteria. Third, as the TALF was closing its final consumer ABS subscription in March of 2010, the ECB announced that the temporary eligibility expansion to non-euro denominated securities would expire in January of 2011. Co-ordination between the two monetary authorities might have been the key to success for both of their programs.

Future research should examine whether these programs had long-term effects in the ABS markets and whether they had any spillover effects. The event study methodology followed in this paper falls short in identifying these long-term effects. The identification assumption in these event studies is that the presence of an announcement is the sole reason why ABS spreads change more than broad market spreads during each short event window. This assumption is reasonable for short-term policy responses, but not for longer event windows; ABS market participants may be reacting to various factors other than one particular program's announcement. Other avenues for future research include the comparison of eligible to non-eligible TALF/ECB program asset classes as suggested by Sundaresan (2011). This study failed to control for such comparisons due to limited data on non-eligible ABS tranches. I note the linear interpolation of missing values for ABS spreads and the aggregation of asset classes in the case of the ECB program evaluation as additional caveats to my methodology due to data limitations.

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NEWLY ADOPTED CORPORATE GOVERNANCE MECHANISM IMPACT ON THE PERFORMANCE OF PUBLIC JAPANESE OVERSEAS ACQUIRERS

Faisal Alzaidi, Aoyama Gakuin University

ABSTRACT

This paper investigates relationships between two main corporate governance components namely the Anti-Takeover Provisions (ATPs) as external component and Ownership Concentration as internal component and the short/long term performance of the Nikkei-listed Japanese cross-border acquirers during the last decade specifically from 2004 to the end of 2013. Market based cumulative abnormal returns (CARs) are used to represent the short term performance. The accounting based metric Return on Assets (ROA) is used to represent long term performance. Based on 222 events, a quantitative methods of events study and regression analysis were employed to reveal the relationships. The study found a negative, weak and statistically not significant relationship between the ATPs and short/long term performance. The other finding is that the relationship between ownership concentration and short/long term performance is almost negligible. These findings imply that the newly adopted corporate governance mechanism in Japan is still not as effective as in other developed markets such as USA and might need more time to reap tangible results.

JEL: G34, G14

KEYWORDS: Anti-Takeover Provisions, Cumulative Abnormal Return, Corporate Governance, Ownership Structure, Events Study, Regression Analysis

INTRODUCTION

The end of the Japanese financial bubbles in early 1990s left the stock market and real estate assets at devastatingly low values. Purchased assets were used by the banks as collateral in the financing process. As these assets values fell sharply the banks ended up in bad positions with regard to these loans and experienced default by many clients. Many Japanese companies lost their source of financing. Domestic demand deteriorated. Excess demand led to a long period of deflation in the local Japanese market. Japan's traditional bank-centered corporate governance system was blamed for the nation's economic bust. The beginning of 21st century marked a notable period of corporate governance reforms in Japan away from bank-centered governance and toward market-oriented governance. Mutual or cross shareholdings among Japanese firms and financial institutions decreased rapidly. At the same time shareholdings by foreign institutions increased for high-performing Japanese firms. Thus, a major divergence in corporate governance occurred among Japanese firms. In the last decade the government of Japan introduced American or Anglo-Saxon style corporate governance after witnessing good performance of Anglo-Saxon based markets such as USA, Canada, UK and Australia. Many reforms related to the corporate governance system were installed to produce a more efficient market. Financial disclosures and transparency were among the targets for improvements.

The last decade also witnessed another notable phenomenon for Japanese firms in the dramatic increase of Cross-Border Merger and Acquisitions (C-B M&A). These mergers were fueled by several factors including: domestic market decline, high cost of local labor and globalization. These factors encouraged Japanese companies to consider entering overseas markets. (C-B M&As) is becoming an important and strategic tool for Japanese companies growth. In 2012, the value of C-B M&As made by Japanese firms hit a new high record of USD 94.5B which accounted for more than 10.5% of the total C-B M&As worldwide (calculated from Bloomberg database). To date there is a lack of research about the relationships between corporate governance and the performance of acquiring firms in Japan, especially firms expanding abroad through acquisitions. Therefore, the aim of this paper is to examine these two remarkable changes affecting Japanese firms in the last decade by investigating:1.) The relationships between the new external corporate governance system represented by Anti-Takeover Provisions (ATPs) and short/long term performance of overseas acquiring public Japanese firms. 2.) The relationship between non-bank-centered ownership concentration representing the new internal corporate governance mechanism and the short/long term performance of overseas acquiring public Japanese firms. The rest of this study is organized as follows: the first section literature reviews of the ATPs and Ownership Structure and their impact on the performance. Section two presents the data, sample construction and research methodology. The next section represents the variables construction for the regression analysis. The last section shows the results with the discussions. The paper closes with some concluding comments.

LITERATURE REVIEW

Anti-Takeover Provisions (ATPs)

The appearance of Anti-Takeover Provisions (ATPs) during 1980s represented the onset of the hostile takeover period. The subject gained the attentions of many practitioners and researchers. ATPs restrict shareholder rights which gives the management the freedom and power to act against any attempt of corporate takeover. The first researchers to address this issue argued about the effect of ATPs on shareholder wealth (DeAngelo and Rice, 1983). In their hypothesis of managerial entrenchment, they argue that ATPs help management protect their positions at the expenses of shareholders and in return reduce shareholder wealth. According to this hypothesis, ATPs increase the agency conflict between managers and shareholders. Many studies seem to support the argument of DeAngelo and Rice that ATPs have negative effect on shareholder value.

Gompers, Ishii, and Metrick (2003), (GIM) examined the effect of ATPs on firm value and shareholder returns. They selected 24 governance provisions incorporated by the Investor Responsibility Research Center (IRRC) and combined these provisions to create a shareholder rights index. They added one point for every provision that works against shareholder rights. Utilizing this index GIM found a significant negative relation between the number of ATPs and firm performance. Low value of the index indicates stronger shareholder rights and the opposite is true. They gave two possible explanations in for the negative relation between stock returns and ATP index. First, investors estimated the true costs of the weak shareholder rights caused by agency problems, which led to share price declines because of these estimates by investors. The other explanation is the classical missing variable explanation. As in corporate governance related studies; some other variables are correlated with the GIM index which causes poorer performance rather than the Index itself. Many later studies tried to explain how the ATPs can influence shareholder wealth and most pf these studies tested the robustness of the GIM finding.

Bebchuk and Cohen (2005) looked at the impact of ATPs on firm value by examining only one specific provision. This provision was staggered boards as a main anti-takeover provision. They found that only a staggered board provision can lead to significantly lower firm value. Bebchuk, Cohen, and Ferrell (2009), extended the results of GIM further by looking at a smaller ATP index based only on six provisions. They argued these provisions were the most important from a legal point of view and have the most influence on firm value. They create a smaller index consisting of 6 provisions from the 24 GIM index and called it the “E index”. They showed that the strong negative relation between GIM Index and firm performance measures is largely driven by the six provision making up the E index. They showed that an index consisting of poison pills, staggered boards, limits to shareholder charter amendments, limits to the amendments of shareholder bylaw, supermajority requirements for large transaction such as mergers and golden parachute have stronger relation and association with firm value and stock returns than the GIM index. This suggests that the other remaining 18 provisions have no significant associations with the firm value. By considering these results the first hypotheses of the research is defined as follows:

H1: ATPs has a negative relationship with performance of the acquiring firm.

Ownership Structure

The other dimension of corporate governance to be investigated in this study is the ownership structure. Specifically, ownership concentration which is an important internal corporate governance component. Earlier researches focusing on the relationship between firm performance and corporate ownership show that firm value increases with ownership of the largest shareholders (McConaughy & Walker, 1998; Claessens &Djankov, 2002; La Porta, Lopez-de-Silanes, Shleifer & Vishny, 2002; Anderson & Reeb, 2003). Andre, Kooli and L`Her (2004) also report that companies owned by large block holders perform better than those owned by smaller investors. The general opinion is the presence of a large shareholder in widely held firms should have a positive impact on firm performance. Agency theory predicts that proper corporate governance mechanisms, such as ownership concentration, can reduce agency problems as stated by (Jensen and Meckling (1976) and Shleifer and Vishny (1986)). The monitoring role of large shareholders can be a beneficial internal mechanism by reducing the agency costs. These shareholders have good incentives and resources to monitor the efficiency level of management and ensure value maximization. McConaughy & Walker (1998) also report a positive relation between concentrated ownership and stock returns.

Two recent meta-analyses studies (Dalton & Daily, 2003; Sanchez and Garcia, 2007) found relationship directions might depend on the institutional environment where the corporation operates. The relation is stronger on the European continent than in Anglo-Saxon countries. This supports the argument that ownership is positively related to firm performance in environments with lower levels of investor protection. Many studies try to identify the effects of institutional ownership on firm performance. Some studies show that institutional shareholders might be good for corporate performance as these institutional shareholders take an active role in corporate governance. But, there is no definite confirmation of the positive impact of such shareholder activism on firm performance (for surveys, see Gillan and Starks, 1998; Black, 1998; Owen, 2005; Romano, 2001). Thomsen and Pedersen (2000) showed empirically that firm performance improves as ownership is more concentrated, but eventually declines in large companies in Europe. This finding implies that, at very high level of concentrated ownership, the positive relationship might turn negative because of the negative effect of the expropriation of small shareholders by large shareholders. LaPorta et al. (1999) found the main problem in large firms might be potential expropriation of smaller shareholders by large controlling shareholders. There are some empirical studies showing that concentrated ownership impact on the performance of acquiring firm is negative. But most empirical research shows a positive relationship. Therefore, the second Hypothesis in this study is built up based on the most prevailing results.

H2: Concentrated ownership is positively related to the performance of the acquiring firm.

The main purpose of this study is to investigate the influence of two main components of the corporate governance mechanism, namely anti-takeover provisions representing the external corporate governance mechanism and ownership concentration representing the internal corporate governance mechanism on the short and long term performance of overseas acquiring public Japanese firms.

DATA AND METHODOLOGY

Sample Construction

The acquisitions sample is extracted from the transaction database of S&P Capital IQ Platform. Some 222 observations are identified between the period January 1st, 2004 and December 31st ,2013. Because two years post acquisitions financial data is needed to gauge long term performance the sample was stopped at the end of 2013. The sample is based on the following criteria: 1.) The acquisition is completed, 2.) The acquirer controls less than 50% prior transaction and majority to 100% after the transaction, 3.) The deal value disclosed is more than $1 million, 4.) The acquirer is a listed public company in the Nikkei225 index which has annual financial statement information available and stock return data (210 trading days prior to acquisition announcements), and 5.) The transaction is cross-border.

Research Methodology

This research employs quantitative methods of data analysis in two steps. The first step entails the event study analysis on the announcement of cross-border merger & acquisitions to determine cumulative abnormal returns (CARs) earned by the acquiring firm's shareholders. The second step is a series of liner multivariate regression analyses to understand the influence of corporate governance namely ATPs and level of ownership concentration on the short and long term performance. To determine short term performance metrics, cumulative abnormal returns with two event windows were used, 2 days before and after announcement date denoted by (CAR2) and five days before and after announcement date denoted by (CAR5). For long term performance measurement, an accounting based metric is used, return on assets (ROA). These three variables were used as dependent variables on the multivariate regression analyses. The main independent variables are Anti-Takeover Provisions Index (ATPINDEX), and Top 5 owner's cumulative percentage (TOP5OW) as the goal of this study is to reveal the influence of theses 2 corporate governance elements on the acquirer's performance. A number of control variables are included as following: Free Cash Flow (FCF), Market Value of Equity (MARKVAL), Leverage (LEVG), Firm size (FIRMSIZE), Relative deal size (DEALSIZE), Cross-border merger & acquisitions. experience (CBMAEXP). The details of each variable can be found in the next section (variables Construction). The following three regression models were used to achieve this study purpose:

$$CAR2 = \alpha + \beta1\ ATPINDEX + \beta2\ TOP5OW + \beta3\ FCF + \beta4\frac{M}{B} + \beta5\ MARKVAL + \beta6\ LEVG + \beta7\ FIRMSIZE + \beta8\ DEALSIZE + \beta9\ CBMAEXP + \varepsilon \quad (1)$$

$$CAR5 = \alpha + \beta1\ ATPINDEX + \beta2\ TOP5OW + \beta3\ FCF + \beta4\frac{M}{B} + \beta5\ MARKVAL + \beta6\ LEVG + \beta7\ FIRMSIZE + \beta8\ DEALSIZE + \beta9\ CBMAEXP + \varepsilon \quad (2)$$

$$ROA = \alpha + \beta1\ ATPINDEX + \beta2\ TOP5OW + \beta3\ FCF + \beta4\frac{M}{B} + \beta5\ MARKVAL + \beta6\ LEVG + \beta7\ FIRMSIZE + \beta8\ DEALSIZE + \beta9\ CBMAEXP + \varepsilon \quad (3)$$

Variables Construction

Acquirer Return

As stated earlier two event windows are used in this study CAR2 is the cumulative abnormal returns 2 days before and after the announcement date e 0, and CAR5 is the cumulative abnormal returns 5 days before and after announcement date. The Cumulative Abnormal Returns is formed by summing individual excess returns over time as in equation (4),

$$CAR_{i,k,l} = \sum_{t=k}^{1} AR_{it} \quad (4)$$

Consistent with Masulis, Wang and Xie (2007) abnormal returns is estimated by using the market model. As shown in equation (5) the difference between the acquirer's stock return (R_{it}) and the expected stock return ($\alpha_i + \beta_i R_{im}$) is estimated with the acquirer's home country as market index. (R_{im}).

$$AR_{it} = R_{it} - \alpha_i - \beta_i R_{im} \quad (5)$$

The market model parameters are estimated utilizing daily stock price data over the 200-day estimation period and in this study is calculated specifically from event day -205 to event day -6.

ROA Variable

Return on Assets (ROA) is used to gauge the long term performance of acquiring firm. By using annual ROA data from the Eikon Thomson Reuters Database, this variable is calculated as the difference between the average value of 2 years post event and the average value of 2 years prior to event.

Anti-Takeover Provisions Index

In this study, an index is formed based on ATP index created by Bebchuk, Cohen, and Ferrell (2009), which known as BCF Index. The BCF index consists of six main components which have the most notable impact on firm performance: staggered boards, limits to shareholder bylaw amendments, limits to shareholder charter amendments, supermajority requirements for mergers, poison pills, and golden parachutes. Since Eikon Thomson Reuters Database has no information on limits to shareholder bylaw amendments and limits to shareholder charter amendments. These components will be substituted by one component which is Significant company transactions (M&A) shareholders' approval component. Therefore, in this study ATP index is based on five components: 1) Poison Pill, 2.) Staggered Boards Structure, 3.) Golden Parachute, 4) Supermajority or qualified majority Vote Requirement, and 5.) Significant Company Transactions (M&A) shareholders' approval. The ATP Index is based on scale from 0 to 5 with higher number representing stronger ATPs undertaken by the firm. Based on the five components, one point is assigned for limiting shareholder rights.

Ownership Structure Variable

This study examines the top 5 owners of the firm based on data available from Eikon Thomson Reuters database and the collective percentage of the top 5 owners during the quarter of which the event (acquisition) occurred and is denoted as (TOP5OW). This variable indicates the concentration level of ownership. Ownership structure is considered as important internal mechanism of corporate governance and it is widely known that it can provide good incentives for large shareholders to effectively monitor management. As the ownership stake of large shareholders or blockholders increases, the greater the incentive to increase firm performance and to monitor closely the management than dispersed shareholders. Also the concentrated actions and monitoring by large shareholders can be easier than diffused or dispersed shareholders. Large shareholders have both the interest and power to get their money back and demand it. There are obvious benefits from concentrated ownership and generally is considered to have positive relationship with firm performance.

Control Variables

Other variables which are of less interest in this study and are controlled for. Free Cash Flow (FCF) is extracted directly from Thomson Reuters Eikon database for the quarter in which the announcement occurred. Jensen's (1986) free cash flow hypothesis argues that FCFs have a negative effect on bidder returns. As managers have more resources available, it becomes easier to engage in empire building. It can however also be argued that higher FCFs are an indication of better firm performance. Performance could be correlated with higher quality managers that tend to make better acquisition decisions. Among the control variables is the Market to Book ratio (M/B), calculated by dividing the Market value of equity by the book value of equity during the quarter which the announcement occurred. This variable represent the growth opportunities. The third control variable is the Market Value of Equity (MARKVAL), defined as the product of multiplying the number of outstanding shares, on the quarter of the announcement, by stock price at the 11^{th} trading day prior to announcement date. Leverage (LEVG) is the fourth control variable. Leverage is often seen as an important governance mechanism. A higher debt to equity ratio reduces future FCFs due to interest obligations and it limits managerial discretion.

Leverage also increases the risk of bankruptcy and provides management with an incentive to improve company performance. Together with debt covenants, managers risk losing control to creditors and might lose their jobs when the firms fall into default. Garvey and Hanka (1999) argue that leverage is related to a firm's takeover protection making it even more relevant as a control variable. Leverage is defined as total debt divided by a firm's market value of total assets during the quarter of the announcement.

The fifth variable is the Firm Size (FIRMSIZE) and is negatively correlated with the acquirers return as shown by Moeller, Schlingemann, and Stulz (2004). They find that on average larger acquirers make acquisitions that generate negative synergies. They also pay a higher premium than smaller acquirers. As in Roll (1986) they interpret the size effect as evidence supporting the managerial hubris hypothesis. Firm size is defined as the natural logarithm of the acquirer's total assets. Relative Deal Size (DEALSIZE) is the sixth variable. It is total transaction value divided by market value of equity during the quarter of the announcement. Moeller, Schlingemann, and Stulz (2004) and Asquith, Bruner, and Mullins (1983) find that acquirer announcement returns increase in relative deal size, but the reverse is true for a subsample of large acquirer in Moeller, Schlingemann, and Stulz (2004). This study focuses only on cross-border Merger & Acquisitions. Experience of acquiring firms in this type of transactions might be a factor which also affect the ultimate performance of the firm. This variable is used a dummy variable with 1 indicating that this is not first time for the firm to go for overseas transactions and 0 indicates that it is first time cross-border acquirer. This variable denoted as (CBMAEXP).

RESULTS AND DISCUSSION

Descriptive and Correlation Statistics

Prior to running the regressions, a descriptive statistics and a bivariate correlation analysis of the dependent and independent variables were conducted. Table 1 presents descriptive statistics for model variables which show that mean and median cumulative abnormal returns, used as metrics for short term performance, are negative and statistically not significant for both two and five-day event windows. Long term performance indicator (ROA) is almost neutral as its mean and median value is 0.14 and 0.2 respectively indicating there is almost no notable changes in the long-run performance. The median value of the ATPs index is 1. We consider any firm with 2 value or above having strong ATPs in place. Firms with 1 value have weak ATPs and firms with zero value have no ATPs. Table 2 shows the correlation matrix. Generally, there is not strong correlation between variables as there is no value over 0.8. The highest correlation value exists between CAR2 and CAR5 which is 0.687. But, both variables are dependent variables and used in separate regression analyses. The next highest value registered in the correlation matrix is 0.664 between firm size and market value of equity. The former represents the book size of the firm and the latter represents the actual and current market size and valuation of the firm. Both variables are used as control variables in the regression analyses which they are of less interest to our study.

Table 1: Descriptive Statistics of the Variables

Variables	Mean	Median	Max	Min	St.D
CAR2	-0.0029	-0.0001	0.0972	-0.1368	0.0395
CAR5	-0.0051	-0.0045	0.1481	-0.2483	0.0549
ROA	0.1459	0.2	16.6	-16.3	4.141
TOP5OW	21.523	19.08	72.31	9.62	10.920
ATPINDEX	1.459	1	3	0	0.8378
FCF	0.0496	0.0353	0.9073	-3.299	0.2980
LEVG	0.9473	0.5225	11.181	0.0006	1.196
M/B	1.476	1.199	12.284	0.167	1.111
FIRMSIZE	27.443	27.494	29.709	24.731	0.8731
MARKVAL	1.604	1.095	8.948	0.0533	1.488
DEALSIZE	0.0008	0.0001	0.029	0	0.0028
CBMAEXP	0.8693	1	1	0	0.3377

This table shows descriptive statistics of the dependent and independent variables. The dependent variables are (ROA) return on assets, (CAR5) cumulative abnormal returns for 5 days before and after announcement day and (CAR2) cumulative abnormal returns for 2 days before and after announcement day. Independent variables are (ATPINDEX) the index for anti-takeover provisions. (DEALSIZE) the relative deal size. (FCF) the free cash flow, (MARKVAL) the market value of equity, (TOP5OW) the aggregate ownership percentage of top5 owners. (M/B) market to book ratio. (FIRMSIZE) the firm size and (LEVG) the leverage measured at announcement day.

Table 2: Correlation Statistics

Variables	ATPINDEX	ROA	CAR5	CAR2	DEALSIZE	FCF	MARKVAL	TOP5OW	M/B	FIRMSIZE	LEVG
ATPINDEX	1										
ROA	-.0049	1									
CAR5	-0.015	-0.131	1								
CAR2	-0.006	-0.091	0.687**	1							
DEALSIZE	0.062	0.053	-0.022	-0.079	1						
FCF	-0.019	-0.005	0.011	0.022	0.027	1					
MARKVAL	-0.315**	-0.033	-0.048	-0.05	-0.122	0.235**	1				
TOP5OW	0.103	-0.059	0.007	0	0.034	0.131	0.146*	1			
M/B	0.087	0.041	-0.170*	-0.075	-0.028	0.06	0.371**	0.189**	1		
FIRMSIZE	-0.365**	-0.023	0.004	-0.084	-0.148*	0.145*	0.664**	-0.163*	-0.155*	1	
LEVG	-0.138*	0.061	-0.064	-0.06	0.064	-0.256**	-0.101	-0.279**	-0.250**	0.162*	1

*This table presents Pearson correlation statistics between the variables of the study (dependent and independent). The dependent variables are (ROA) return on assets, (CAR5) cumulative abnormal returns for 5 days before and after announcement day and (CAR2) cumulative abnormal returns for 2 days before and after announcement day. Independent variables are (ATPINDEX) the index for anti-takeover provisions. (DEALSIZE) the relative deal size. (FCF) the free cash flow, (MARKVAL) the market value of equity, (TOP5OW) the aggregate ownership percentage of top5 owners. (M/B) market to book ratio. (FIRMSIZE) the firm size and (LEVG) the leverage measured at announcement day. ** and * indicate significance at the 5 and 10 percent levels respectively*

Regression Analysis Results

Table 3 presents the results of the three regression analysis carried out in this study. The results show that the Anti-Takeover Provisions (ATPs) as an external corporate governance mechanism has a negative but statistically non-significant effect on the short term performance of the acquirer. This is represented by the two indicators, cumulative abnormal returns with 2 days before and after the announcement day (CAR2) and 5 days before and after the announcement day (CAR5). ATPs also has a negative influence on long term performance represented by (ROA) as shown on regression C. But ROA is statistically not significant. Generally speaking, ATPs have only minor negative influence on the short and long term performance of overseas acquiring public Japanese firms. This finding support weakly the hypothesis H_1: Anti-Takeover provisions have a negative relationship with the performance of the acquiring firms. This is weak support because all the three regression analyses revealed statistically non-significant relationships between ATPs and short-term performance as well as long term performance of acquiring firms. This appears inconsistent with the strong negative association documented in Masulis, Wang and Xie (2007). However, Core, Guay, and Rusticus (2006) and Bebchuk, Cohen and Wang (2013) argue that the adverse impact of ATPs has been positively moderated in the period after 2001. Since our sample includes more recent acquisitions than those used by Masulis, Wang and Xie (2007), our results are likely to reflect the diminishing association between ATPs and firm performance. We find statistically non-significant and very weak relationships (Beta = 0) found also between the internal corporate governance component represented through ownership concentrations (TOP5OW) and short term performance of acquiring public Japanese firms represented by CAR2 and CAR5. The relationship become negative with long term performance indicator (ROA) but statistically insignificant. Therefore, this finding of insignificant relationship between the ownership concentration and both short and long term performance of the acquiring public Japanese firms does not support our second hypothesis

H2: The concentrated ownership is positively related to the performance of the acquiring firm.

The Coefficient of determination (R2) values for the three regression models explains clearly the insignificant effect of the independent variables specifically ATPs and Ownership concentration on the dependent variables CAR2, CAR5 and ROA which are the indicators of firm short and long term performance. From R2 values, dependent variables only explain 3.6%, 4.3% and 1.9% of the variations of the dependent variables of regressions A, B and C respectively. Only M/B ratio variable has a statistically significant relationship with CAR5 at the 5% level but this variable is a control variable and not of interest to this study.

Table 3: Regression Results

	Regression A (Dependent: CAR2)		Regression B (Dependent: CAR5)		Regression C (Dependent: ROA)	
	β	P (Sig.)	β	P (Sig.)	β	P (Sig.)
ITERCEPT	0.291	0.042	0.125	0.546	-2.189	0.890
ATPINDEX	-0.001	0.710	-0.001	0.865	-0.357	0.336
TOP5OW	0.000	0.639	0.000	0.912	-0.018	0.533
FCF	0.003	0.719	-0.002	0.906	0.368	0.716
M/B	-0.006	0.071	-0.011	0.012**	0.410	0.233
MARKVAL	0.004	0.254	0.002	0.632	-0.265	0.461
LEVG	-0.002	0.469	-0.005	0.157	0.200	0.452
FIRMSIZE	-0.010	0.059	-0.004	0.584	0.108	0.851
DEALSIZE	-1.313	0.168	-0.448	0.733	71.214	0.478
CBMAEXP	0.001	0.881	0.003	0.818	-0.203	0.823
R2	0.036		0.043		0.019	

*This table presents results of the three regression analyses: regression A with (CAR2) the cumulative abnormal returns for 2 days before and after announcement date as dependent variable, regression B with (CAR5) the cumulative abnormal returns for 5 days before and after announcement date as dependent variable and regression C with (ROA) return of assets as dependent variable. Independent variables of interest are (ATPINDEX) the index for anti-takeover provisions and (TOP5OW) the aggregate ownership percentage of top5 owners. Other independent variables are (DEALSIZE) the relative deal size. (FCF) the free cash flow, (MARKVAL) the market value of equity, (M/B) market to book ratio, (FIRMSIZE) the firm size and (LEVG) the leverage measured at announcement day. R2 represents the coefficient of determination. ** indicate significance at the 5 percent level.*

CONCLUDING COMMENTS

Utilizing 222 observations, this study focuses on the influence of two aspects of corporate governance mechanism. Specifically we focus on Anti-Takeover Provisions as an external corporate governance element and Ownership Concentration as an internal corporate governance element on the short/long term performance of the cross-border public Japanese acquirers in the last decade. Quantitative methods of data analysis were used. Specifically, an events study and regression analysis are used to achieve the goals of this study. The event study methodology was used to calculate two short term market based metrics namely cumulative abnormal returns 2 days before and after the event (CAR2) and 5 days before and after the event (CAR5). For a long term performance indicator, an accounting based metric, Return on Assets (ROA) was used. The findings reveal that ATPs has almost a neutral influence on the short and long term performance of overseas acquiring public Japanese firms in the last decade. This finding does not support the managerial entrenchment hypothesis. Moreover, we cannot find strong evidence for self-dealing as there is no real shareholder value destruction caused by the cross-border merger & acquisitions. Performance did not change in the short or long run.

It seems that investors should be less concerned about ATPs in Japan as it seems that they are not as effective as in other developed markets. This study also finds no influence of ownership concentration, represented by the cumulative percentage of the top 5 owners, on the short and long term performance of cross-border acquirers. Almost no relationship exists, implying that an institutional and foreign ownership system, which replaced the old bank centered system in the last decade, still have not worked to enhance shareholder value. The limitation of this study is that we use only one Anti-Takeover Provisions Index. There were not many corporate governances related data sources available to construct several indices which can be used for comparison purpose and for testing the robustness of the index. For future research, more accounting based metrics such as Tobin Q and ROE can be included as dependent variables. The long term market based event study methodology, specifically the Buy and Hold Abnormal Returns (BHAR) is recommended to confirm the current study findings and to have better understanding.

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TESTING FOR STOCK PRICE BUBBLES: A REVIEW OF ECONOMETRIC TOOLS

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ABSTRACT

This paper presents an overview of several econometric tools available to test for the presences of asset price bubbles. For demonstrative purpose, the tools were applied to historical stock price and dividend data starting from 1871 through 2014. The earliest tools developed were Shiller's variance bound tests and West's two step procedure. Though these tools are useful in detecting asset prices, they are subject to some serious econometric issues. To address these limitations, Cointegration methods were used to detect asset price bubbles. Unfortunately, if there are collapsing bubbles, Cointegration techniques cannot identify multiple bubbles. To overcome this Phillips, Shi and Yu (2015) developed a right tailed Augmented Dickey-Fuller test. This test not only identifies multiple bubbles but also dates the starting and ending period of a bubble. Availability of such real time monitoring tool would significantly help investors, retirees, and portfolio managers to rebalance their portfolios during such bubble periods.

JEL: G12, G14

KEYWORDS: Stock Price Bubble, Cointegration, and Right Tail ADF

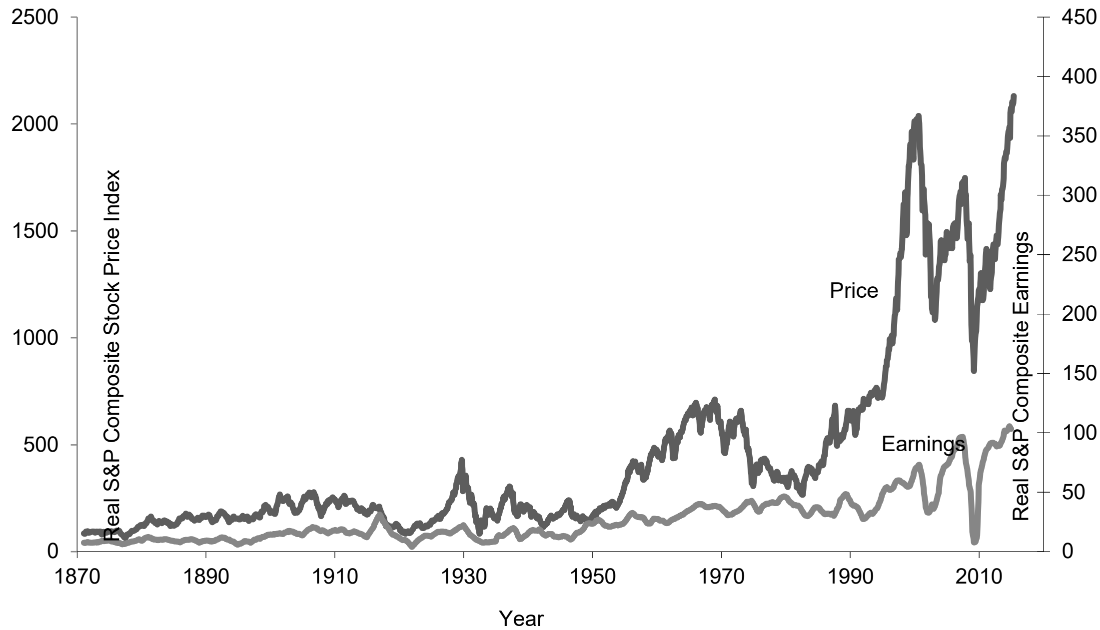
INTRODUCTION

A bubble in asset prices occurs when the present value models persistently fail to explain asset price levels greatly exceeding the value justified by fundamentals. The Finance profession is of two minds on the possibility of bubbles. On one hand the Efficient Markets School argues that asset prices are determined by the available information and this precludes the possibility of financial bubbles. They argue that when market prices deviate from the underlying fundamental values, arbitrage will force market prices to align with fundamental values. On the other hand, the Behavioral School rejects that asset prices are solely determined solely by the present value relationships. They believe investors are subject to a host of psychological biases and irrational impulses and their decisions to buy or sell may lead asset prices to deviate from their fundamental values. At the theoretical level they offer several explanations. For example, behaviorists note investors are psychologically primed to forecast current trends to continue (extrapolation bias). So that current price rises are predicted to continue to rise at the current rate. Alternatively herding behavior inclines investors to place their money in the direction of current market trends. Therefore, they conclude that in addition to fundamental factors, other psychological factors play a role in the determination of asset prices. Such theoretical disputes are best settled by examining the empirical evidence. For example, stocks prices are expected to reflect discounted future earnings.

Figure 1 shows the monthly real earnings and the real S&P 500 stock prices from 1871- 2014. While earnings were fairly stable over the whole period the S&P 500 stock prices spurted modestly in the late 1960s and early 1970s but were closely aligned with real earnings. However, beginning early 1980s stock prices rose sharply particularly relative to earnings. The sharp rise in stock prices (particularly relative earnings) in about 1997 and the subsequent decline at the end of the century led the general public and financial press to characterize the period as "the dotcom" or "the internet bubble." The upward spike in

2007 and the subsequent stock price collapse portrayed in Figure 1 led the general and financial press to characterize the period of 2007-2010 as the "Real Estate bubble". The press attributed the recession at that time to the bursting of the bubble. The recent rise in stock prices has led the press and even academics to speculate that another bubble is forming. For example, Robert Shiller in a recent comment while expressing uncertainty claims "there is a bubble element to what we see." (www.businessinsider.com/robert-Shiller-stock-market-bubble-2015-5).

Figure 1: S & P Price Index and S&P 500 Composite Earnings Sample: Monthly Data January 1870 - December 2014



This figure shows the monthly values of the real S&P 500 price index and the real S&P 500 composite earnings for January 1870 through December 2014. The data source is Robert Shiller's website.

In recent years a number of econometric tools have been developed to test for financial bubbles. These tests have significant practical implications. If bubbles are present, investors, retirees, portfolio managers, regulators and policy makers, would like to detect them in order to take appropriate countermeasures. Investors, retirees, and portfolio managers would seek to rebalance their portfolios while the regulators and the policy makers could adopt appropriate policies to limit the damage to the real economy. This paper examines the econometric techniques available to detect the presence of bubbles. The original techniques focused on identifying the presence of a single financial bubble. The latter techniques enhanced our ability to spot a single bubble and in addition provided the capability to discern multiple financial bubbles and to recognize their beginning and ending points.

In the next section we review some of the previous literature and present descriptive statistics on all the variables employed in our econometric tests. We explain the methodology for the three types of bubble detection tests we utilize. The first methodology we explain is the Variance Bound Test. This test is representative of tests that investigate the consistency of elevated stock prices with the present value of the dividends model. The second test methodology examined is cointegration tests. These tests spot bubbles by studying the time series properties of stock prices and dividends. We then explain the methodology of another more sophisticated time series test which is capable of detecting multiple bubbles and to date the beginning and end of a bubble. In the following section we report the results of these three tests. In the closing section we offer our conclusions.

LITERATURE REVIEW

Shiller (1981) and LeRoy and Porter (1981) first developed variance bound to monitor the present value models under the assumption of rational bubbles. Although the purpose of these tests was to evaluate the present value of dividends model, Blanchard and Watson (1982) and Tirole (1985) among others suggested that a rejection of present value of dividend model is consistent with a bubble. Thus, the test may be used to substantiate the presence of a bubble.

Although the Variance bounds test is one of the first options developed for testing and identifying financial bubbles, we do not perform it in this paper. This is because the test has serious problems. First it tests a joint hypothesis. The test simultaneously rejects the Present Value model and thereby is unable to reject the presence of a bubble. Further, Kleidon (1986) shows the test breaks down if the data is non-stationary. Gurkaynak (2008) provides a more detailed discussion of this and other Variance bound test issues. A variation of Variance bounds test was proposed by West (1987) which can be adopted even if the data is non-stationary. Diba and Grossman (1988) note that the present value of dividend model does not allow for the possibility of a bubble starting. Thus a bubble is most likely the result of some unobserved variables. To uncover these unobserved fundamentals in the time series properties of the data Campbell and Shiller (1989) and Diba and Grossman (1988) adopted unit root and cointegration tests, respectively, to detect asset price bubbles. We will present a version of this test later in the paper.

Evans (1991) pointed out that these tests suffer from a serious limitation. He argues that these techniques cannot detect the presence of multiple bubbles in a long time series. To overcome this limitation, Phillips et al (2011) proposed a right tailed Dickey-Fuller test for detecting and dating asset price bubble. Phillips et al (2105) then generalized the right tailed Dickey-Fuller to identify and date multiple bubbles in a long time series data. We first focus on testing and dating the presence of a stock price bubble in the US. However, since US stock markets have gone through several bubbles, we also employed the generalized right tailed Dickey-Fuller tests proposed by Phillips et al (2015) to detect presence of and to date multiple bubbles. We will pay close attention to this test later in the paper.

DATA AND METHODOLOGY

The real S&P 500 total annual returns, real prices, real earnings and the real annual dividends were obtained from Robert Shiller's website for the period 1871-2014. The annual data are used to test for the presence of a bubble but monthly data from 1960 through 2014 were used to identify and date multiple bubbles. Table 1 provides descriptive statistics of annual data for the real S&P 500 prices and earnings for the periods 1871-2014 and for the end of the period 1960-2014. The mean level of Real S&P 500 Prices is substantially higher (about 95% increase) in the 1960-2014 period compared to the whole period, 1871-2014 (796.99 vs.409.26). In contrast, real Index dividends (13.17 vs. 20.06) only increased by about 52%. Obviously there is a clear miss-alignment between prices and dividends.

Table 1: Descriptive Statistics of Annual Data

	Real S&P 500 Index Price	Real Index Dividends	Real S&P 500 Earnings	Ratio of Real Price to Real Dividends
Panel A: Sample 1871 - 2014				
Mean	409.26	13.17	25.06	26.83
Median	232.87	11.63	18.33	22.09
Maximum	1975.56	31.45	92.73	90.89
Minimum	75.89	4.68	3.95	12.41
Std. Dev.	412.16	6.15	18.23	13.92
Skewness	2.00	0.72	1.52	2.23
Kurtosis	6.36	2.76	5.27	8.70
Jarque-Bera	160.01	12.36	84.47	314.55
Probability	0.00	0.00	0.00	0.00
Panel B: Sample 1960 - 2014				
Mean	796.99	20.06	43.76	37.81
Median	612.33	20.02	37.55	32.30
Maximum	1975.56	31.45	92.73	90.89
Minimum	290.92	15.31	24.49	17.70
Std. Dev.	463.23	3.83	16.90	16.97
Skewness	1.01	1.18	1.35	1.40
Kurtosis	2.72	4.10	3.96	4.55
Jarque-Bera	8.94	14.64	17.82	22.33
Probability	0.01	0.00	0.00	0.00

This Table provides descriptive statistics for each of the variables discussed in our study. It covers the entire sample period as well as the modern period. The data used are annual data.

Variance Bound Tests

As discussed previously one of the first bubble detection procedures was originated by Shiller (1981) and LeRoy and Porter (1981). We noted the reasons we do will not replicate this test. We shall employ a similar and related procedure, West's (1987) two-step test. This test is designed to overcome the objections to the variance bound test and represents a significant advancement in bubble detection. This test avoids the problem of joint hypothesis testing by directly testing for the null hypothesis of no bubble. In addition, the test is valid even if the prices and dividends are non-stationary. To understand West's test, it is useful to present a two period version of Shiller's model. The present value of dividends model for the two period case is:

$$P_o = \frac{E(D_1)}{1+r} + \frac{E(P_1)}{1+r} \tag{1}$$

Where P_o is the current price, $E(D_1)$ is the expected dividend in year 1 and $E(P_1)$ is the expected price in year 1 and r is the real expected rate of return on the stock market. To apply this formula, the problem we face is that we don't know the expected dividend stream at time 1 nor the terminal price. Shiller (1981) cleverly finessed this problem. He notes we know the past dividends and also assumes the discount rate is known and constant. Suppose we have 100 years of dividends and stock price data, then we can find a perfect forecast value at time 1 of the 100 years, p^*_1. It is the present value of dividends years 2 through 100. (This is an approximation as it neglects dividends beyond 100. This error should be small if we truncate the sample and assign a terminal value which is average stock price of the entire sample). The perfect forecast value and the actual price differ only by the forecast error.

$$P_t^* = P_t + \varepsilon_t \tag{2}$$

If rational expectations hold then $\mathcal{E}_t$ and P_t are independent. Then equation (2) implies

$$Var(P^*) = Var(P_t) + Var(\varepsilon_t) + Cov(P^*, \varepsilon_t) \quad (3)$$

and the covariance term is zero. Thus, Var(p*) ≥ Var (P_t).

So if the variance of the actual stock prices exceeds the variance of the perfect forecast prices then the dividend discount model does not hold.

The dividend discount model is

$$P_t = \gamma E(P_{t+1} + d_{t+1}/I_t) \quad (4)$$

where I_t is the information available to the investor at time t and ϒ is the present value interest factor for the discount rate, r. That is ϒ=1/(1+ r_t). It may be computed in a regression format with observable variables.

$$P_t = \gamma(P_{t+1} + d_{t+1}) + U_{t+1} \quad (5)$$

Where $U_{t+1} = -E(P_{t+1} + d_{t+1}) - \gamma(P_{t+1} + d_{t+1})$

West (1987) using the dividend discount model as basis, provides a two- step procedure to test for the presence of a bubble. Instead of using the Ordinary Least Squares regression, West employs an instrumental variables approach utilizing dividends as the instruments. West (1987) assumes that the dividends follow a first order autoregressive (AR (1)) process where

$$d_t = \beta d_{t-1} + V_t \quad (6)$$

He shows this implies

$$P_t = (\frac{\gamma\beta}{1-\gamma\beta})d_t + \varepsilon_t \quad (7)$$

There are two ways to estimate $(\frac{\gamma\beta}{1-\gamma\beta})$ in equation (7). We can use the direct result of equation (7) or the indirect result of equations (5) and (6). In the absence of a bubble the two methods will produce the same results. The null hypothesis of no bubble will not be rejected. To perform the test, we use our estimate of γ from equation (5) and the estimate of β of equation (6) and compute $(\frac{\gamma\beta}{1-\gamma\beta})$. If this value lies in the 95% confidence interval of our estimate of the coefficient, $(\frac{\gamma\beta}{1-\gamma\beta})$, from equation (7) we do not reject the null hypothesis of no bubble. Otherwise it is rejected and we conclude the data supports the presence of a bubble. The real S&P 500 total annual return and the real annual dividends used in our test were obtained from Robert Shiller's website for the period 1871-2014

Cointegration Tests

Both the variance bound test and West's test attempt to uncover significant deviations from a rational stock valuation model. Diba and Grossman (1988) advocate examining the time series properties of stock prices and dividends and the stability of the underlying relationship between them. Thus it is logical to ascertain whether stock prices are stationary or how much differencing is required to impose stationarity. If stock prices are more explosive over time than dividends, then this may be attributed to an asset price bubble. Thus, it is natural to test for the presence of a bubble by examining stock prices and testing for unit roots.

If we find that stock prices have unit roots while earnings do not, then it can be construed as evidence against the presence of a bubble. However, if both stock prices and earnings have unit roots then it is still necessary to probe the underlying stability of the relationship between them. One way of testing for stability in their underlying relationship is by using a statistical procedure called Cointegration. Two variables are said to be cointegrated if they share a common long run stochastic trend. However, if the relationship between earnings and stock prices becomes unstable during a period then that would support the presence of a bubble. Cointegration methodology is well suited to test for the presence of asset price bubbles. Before conducting formal tests, it is useful to review Figure 1 of real stock prices and real earnings. The plot shows that before 1982 real stock prices and real earnings appear to move together. After 1982 stock prices increased far more explosively than earnings. This raises the possibility that a cointegration test may reveal a break in the nexus between stock prices and earnings. The first step to establish cointegration is to test for the presence of unit roots. Several econometric approaches were developed in the literature to test for unit roots. For demonstrative purposes here we used the Augmented Dickey-Fuller (ADF) test to verify the presence of a unit roots in both stock prices and dividends. The second step entails testing for cointegration. To test for cointegration we used the Johansen-Juselius procedure (1988). This is accomplished by estimating the following cointegration equation:

$$P_t = a + bd_t + U_t \quad (8)$$

Where for our 2 variable case P_t is the S&P 500 Index and d_t is the dividends on the index, an and b are regression parameters and U_t is the random error term. We may rewrite the equation as

$$\Delta P_t = a + (b-1)d_{t-1} + e_t \quad (9)$$

Where e_t is the error term of equation (9). If the variables are cointegrated then b, the slope term, will be equal to one and the null hypothesis of no cointegration is rejected. On the other hand, if b is close to zero then we will be unable to reject the null of no Cointegration. The Johansen-Juselius procedure views this equation in a matrix format. The presence of cointegration is determined by the rank of b matrix. The highest rank of b that can be obtained is n, the number of variables under consideration. If b is zero, there are no linear combinations that are stationary and so there are no cointegrating vectors. In the two variable case we consider, the matrix is a 1x1 consisting of the b-1 term. The rank can only be zero or one. The rank of the matrix is determined by the magnitude of the eigenvalue of the matrix. The rank of the matrix equals the number of nonzero eigenvalues. If the eigenvalue is not statistically different from zero, then the rank is zero. We do not reject the null hypothesis ($b \neq 1$); this leads to the conclusion that the variables lack Cointegration. If the eigenvalue is statistically different from zero, then we conclude the rank of the matrix equals one. This is consistent with cointegrated variables. To test this, we use the annual S&P 500 prices and the real dividend index for the period 1871-2014 from Shiller's website.

Time Dating and Multiple Bubble Tests

A method developed by Phillips et al (2011) addresses the issue of time dating financial bubbles. It constructs a right-tail Dickey-fuller test to identify the start and the end date of a bubble and provides greater power than the cointegration methodology. Still the test is unable to identify multiple bubbles. Further, the presence of multiple bubbles usually lowers the power of this testing methodology. This also applies to the West's test and Cointegration tests. This adds increased importance to the search for statistical methods capable of dealing with multiple bubbles. The Methodology of Phillips et al (2015) generalizes the analysis of Phillips et al (2011) to identify the start and end points of multiple bubbles. Both methods are based on the following reduced form equation:

$$P_t = \mu + \delta P_{t-1} + \sum \varphi_i \, \Delta P_{t-1} + e_t \quad (10)$$

where P_t is the price of the S&P 500, μ is the intercept, and e_t is the random error term. The δ is the key coefficient estimated by the regression. This ADF statistic (t-statistic) for this regression is not quite the same as the standard unit root ADF tests. This is because the standard unit root tests are left tail tests and while the unit root test proposed by Phillips et al (2015) are right tailed tests. We need the right tail, not the left tail values of this asymmetric distribution. Thus the critical values of the right tailed ADF statistic will differ from those used in the standard left tailed unit tests. The null hypothesis is the data contains a unit root and the alternative hypothesis postulates the presence of a mildly explosive autoregressive coefficient. Formally, the null and alternative hypotheses are:

H_o: δ =1
H_1: δ >1

Next we explain the test procedure in terms of the specifics of our dataset rather than present it in a more generalized form. With annual data it is difficult to identify multiple bubbles with periodically collapsing behavior and to precisely date when the bubble started and when it ended. Thus, to detect multiple bubbles it is important to have higher frequency data and therefore we moved from annual to monthly data. Our dataset consists of 660 monthly S & P 500 real price to dividend ratio observations covering the period of 1960 through December of 2014. We shall sample the data in fixed window sizes of 53 monthly observations, thus the first sample covers observations 1- 52. For this interval an ADF statistic is computed. Then the interval is then increased by one unit (1- 53) and an ADF_{53} is calculated. This continues until the interval covers the whole sample (1, 660). Thus 607 (660-53) ADF statistics are computed. The SADF statistic is the Supremum value of the set of ADFs calculated. A bubble occurs if the ADF exceeds the critical value of the statistic. Critical values of the SADF are set by Brownian motion of stock price movements. Critical values of the ADF statistic for each date are derived from the distribution of the ADF statistic by Monte Carlo Methods. The details may be found in Phillips et al (2011).

The Generalized SADF (GSADF) offers greater statistical power than the SADF statistic and also adds the capability of dating multiple bubbles. The first step of the GSADF test is to perform the same recursive regression as the SADF test. However, in this test instead of fixing the starting point of the recursion on the first observation, GSADF changes both the starting and ending points of the recursion. For example, a starting point of the recursion could be from 10th observation through 63rd observation. Then add one period at a time running a regression for each period until the final regression (660-53) period 607 to period 660. The GSADF for a given date is the supremum value of all the GSADFs calculated with an interval ending on that date. Since GSADF test includes more sub-samples with a flexible starting and ending windows, it does a better job of identifying multiple bubbles in the data. The distribution of the GSADF, like the SADF, are set by Brownian motion of stock price movements. This allows the use of Monte Carlo methods to find the critical values. Again Phillips et al (2015) provide details.

The methodology of dating the bubbles in real time for the GSADF test involves performing a backward SADF (BSADF) test on an expanding sample sequence. To make this concrete suppose we are considering a total sample of 200 periods. The first BSADF statistic (backwards ADF) is calculated from (200-53) 147 to 200. The second covers 146 to 200 and so on until 0 to 200. This is a special case wherein of course, the test would need to be performed on each period with changing starting and ending points. A bubble is identified when a BSADF crosses the 95% critical value from below and terminates when it crosses it from above.

RESULTS

To demonstrate the implementation of West's test we used the US stock market data. Again, the real S&P 500 total annual return and the real annual dividends were obtained from Robert Shiller's website for the period 1871-2014. The real S&P 500 price series displayed in Table 1 was discussed previously. The dividend series is also shown in table 1. The dividends were considerably more stable than the real S&P 500 price time series as evidenced by higher standard deviation in price series in the whole period (412.16 versus 6.15). The test was conducted for the whole sample period as well as the recent period (1982-2014) where earnings and the stock prices diverged. As we discussed previously the West Test is based on a comparison of the results of a direct estimate of γ as shown in equation 7 with the results of the indirect estimate of γ and β from equations 5 and 6. The results of this are reported in Table 2. The results show the estimated coefficients of γ and β for the whole as well as for the sub sample period. Surprisingly the coefficients are stable and statistically significant for the whole and the sub-sample period. The results of estimated coefficient of equation (7) is presented in Panel C of Table 1. The coefficient is statistically significant in both periods.

Table 2: Regression Results for Equations 5 Through 7 for West Test Annual Data

Panel A		
	Equation 5: $P_t = \gamma(P_{t+1} + d_{t+1}) + U_{t+1}$	
	WHOLE PERIOD 1871-2014	1982-2014
	EQUATION 4	EQUATION 4
Coefficient	γ = .9359	γ =0.9349
Std Error	0.0145	0.282
T-Stat	64.33***	23.11***
P-Value	0.000	0.0000
Panel B		
	Equation 6: $d_t = \beta d_{t-1} + V_t$	
	WHOLE PERIOD 1871-2014	1982-2014
Coefficient	β = 1.0194	β = 1.0328
Std Error	0.0079	0.0145
T-Stat	129.41***	71.27***
P-Value	0.000	0.000
	Equation 7: $P_t = (\frac{\gamma\beta}{1-\gamma\beta})d_t + \varepsilon_t$	
Coefficient	$(\frac{\gamma\beta}{1-\gamma\beta})$ = 23.360	$(\frac{\gamma\beta}{1-\gamma\beta})$ = 38.70
Std Error	1.721	2.511
T-Stat	13.58***	15.41***
P-Value	0.000	0.000

*This table estimates the regressions necessary to perform the West Test. Equation 4 is a statistical representation of the present value of the dividends model. Equation 5 is an AR (1) representation of the dividends series. Finally, equation 7 combines equations 5 and 6. ***statistically significant at 1 percent*

Table 3 reports the results of testing the null hypothesis of no bubble. If there is no asset price bubble the direct estimate of the parameter in equation (7) should equal the regression estimates of equation (5) and (6) plugged into the formula $(\frac{\gamma\beta}{1-\gamma\beta})$. The γ for the whole period is 0.9359 (Equation 5) and for β is 1.019. This implies an indirect estimate of the coefficient of equation (7) is 20.7865 (9359*1.0194/(1-0.9359*1.0194)). The direct estimate shown in Table 2 for equation 7 is 23.36, its standard error is 1.721. This leads to a 95% confidence interval of (19.92, 26.80). Thus the confidence interval captures the value of direct estimate of equation (7) so the test fails to reject the null hypothesis of no bubbles.

Table 3: West Tests for Financial Bubbles

Annual Data Total Sample: 1871-2014						
Ho: No Financial Bubble **Total Sample: 1871-2014** Period	γ	β	Indirect $(\frac{\gamma\beta}{1-\gamma\beta})$	Direct $(\frac{\gamma\beta}{1-\gamma\beta})$	95% CI	Decision
1871-2014	0.9359	1.019	20.79	23.36	(19.92, 26.80)	Do Not Reject
1983-2014	0.9349	1.033	28.04	38.70	(33.68, 43.73)	Reject

This table reports the results of the Two Step West Test for the whole period (1871-2014) and the more recent period (1983-2014). The second and third columns reports the estimates of γ and β obtained from equations 5 and 6 in Table 3 for both periods. Column 4 uses these results used to indirectly estimate the coefficient of equation 6 (for both periods). Column 5 reports the direct estimates of equation 6. If the null hypothesis of no financial bubble holds, the estimates should be equal the indirect estimates on column 5 and should be captured in the 95% confidence interval of equation 7 shown column 6.

We are most interested in the possibility of a bubble in a relatively recent period. Information from the 19th century or early or even mid-20th century probably offers little aid in detecting bubbles in the more recent period. Thus we considered the period when stock prices increased more than earnings (1982-2014). It is recent enough to shed light on the current situation and long enough to provide an adequate number of observations. The γ for the sub-sample period is 0.9339 (Equation 5) while the estimate of β is 1.0328. This implies an indirect estimate of the coefficient of equation 7, $\frac{\gamma\beta}{1-\gamma\beta})$ of 28.04 (0.9349*1.0328/(1-0.9349*1.0328)). The direct estimate shown in Table 2 for equation 7 is 38.7041, with a standard error of 2.51141. This leads to a 95% confidence interval of (33.68, 43.73). Thus the confidence interval does not include the indirect estimate and the test rejects the null hypothesis of no bubbles. The West Test is consistent with the presence of a bubble or bubbles during this time period. The West test failed to detect a bubble in the whole time period. Thus our results suggest the necessity to place close attention to sub periods as the whole period may mask variation. It does not indicate when the bubble originated or when it burst. Even so, the West Test only detects bubbles but does not indicate the time of origin or the end time. Cointegration tests offer a way to pay close attention to the statistical relationship between real stock prices and real dividends during the whole period and more importantly during sub periods. If a long term equilibrium relationship between real stock prices and real dividends were to break down during a period of stock price increases, this would be consistent with the presence of a bubble. The test for cointegration ultimately tests for the presence of a common stochastic trend. If variables real stock prices and real dividends are cointegrated prior to a stock price run-up but are no longer cointegrated during the run-up, then we might infer the presence of a bubble.

Before testing for Cointegration, the stationarity of the variables must be determined. Figure 1 is clearly consistent with nonstationarity for stock prices. Still it is useful to perform more formal tests. Tests for stationarity are performed using the augmented Dickey-fuller (ADF) tests. The test allows for a drift term, a linear time trend, and for multiple period lags. The null hypothesis is that the variables are non-stationary; the alternative hypothesis is that they are stationary. The critical values for the ADF test are reported in Engle and Yoo (1987). The results of these stationarity tests are shown in Table 4. The lag length was determined by the Bayesian criterion. For both real stock prices and real dividends, the ADF test supports nonstationarity at a one per cent level of significance.

Table 4: Augmented Dickey-Fuller Unit Root Test

	Sample: Annual 1871-2014 13 Period Lag	**Sample: Annual 1983-2014 1 Period Lag**
Variable	**Augmented Dickey Fuller Test Statistic**	**Augmented Dickey Fuller Test Statistic**
Real S&P 500 Stock Prices	8.84 ***	-1.09***
Real S&P 500 Dividends	1.01***	1.94***
Critical Value (1%)	-3.48	-3.65

*This table presents the results of a formal test of the stationarity of the Real S&P500 Prices and Real S&P 500 dividends. The null hypothesis is the time series are stationary. ***Significant at 1 percent level*

Table 5 presents the results of these cointegration tests for the whole period, 1871-2014 and the 1983-2014 periods. For the whole period the Trace Statistic Test shows that for we may reject at a five per cent level of significance the null hypothesis that no eigenvalue is different from zero. Thus we may reject the null hypothesis of no cointegration at the 5% level of statistical significance. Thus real stock prices and real dividends are strongly linked. For the later portion of the sample 1983-2014 the trace tests indicate that eigenvalues are not statistically distinguishable from zero. Thus, the test fails to reject the null hypothesis that stock prices and dividends lack cointegration during this time period. This result suggests that the linkage between real S&P 500 prices and real dividends has been substantially reduced during the period of 1983-2014. The statistical evidence is consistent with the presence of a financial bubble.

Table 5: Cointegration Tests for the Presence of Financial Bubbles

Cointegration Between	**Hypothesized No. of CE(s)**	**Eigenvalue**	**Trace Statistic**	**0.05 Critical Value**	**Probability**
Real S&P 500 Stock Prices vs. Real Dividends annual 1876-2014	None** at most 1**	0.082 0.038	17.4 5.5	15.49 3.84	0.027 0.02
Real S&P 500 Stock Prices vs. Real Dividends annual 1983-2014	None at most 1	0.215 0.000	7.76 0.005	15.49 3.84	0.49 0.94

*For the time period 1876-2014: Trace test indicates 2 cointegrating equations and the null hypothesis of no cointegration is not rejected. For the time period 1983 -2014: Trace test is unable to reject the null hypothesis of no cointegration. **significant at 5 percent level.*

Both West's test and Cointegration successfully detected the presence of a bubble in the 1983-2014 period. The Cointegration test offers some improvement over the West Test in dating the bubble. The West Test only detects a bubble in the sample period and offers no guidance for the ascertaining the start of the bubble. Cointegration tests suggest an approximate starting point. The bubble starts when the cointegration relationship breaks down. Still we cannot precisely pinpoint when that occurs. Further, as we have already suggested, these methods suffer from two shortcomings. First as demonstrated by Evans (1991) these tests will detect only permanent bubbles. They fail to detect bubbles that collapse and perhaps even restart. Van Norden and Vigfusson (1998) and Hall, Psaradakis, and Sola (1999) attempted to remedy this deficiency by treating expanding and collapsing bubble as different regimes in a Markov process. Still this method cannot distinguish between a single and multiple bubbles. This is particularly important for our purposes. Both methods identified the presence of a financial bubble during 1983-2014 period. Neither method attributed the bubble to the internet, or real estate boom or both. The rise of the S&P 500 from a low of 736 to 2122 in June 2015 led the financial press to conjecture that a second equity bubble has begun. Neither of these techniques can address this issue. The SADF and GADF tests developed by Phillips et al. (2011) and Phillips et al (2014) attempts to deal with these difficulties. These tests offer the advantage of time dating the beginning and the end of a bubble. They also are capable of detecting multiple bubbles. The SADF Test establishes the start of a bubble as the first date the ADF series crosses the critical value series from below and the end of the bubble when the ADF series crosses the critical value series from above. The estimated SADF statistic for the period and the critical values for 90%, 95%, and 99% confidences respectively are presented in Table 6.

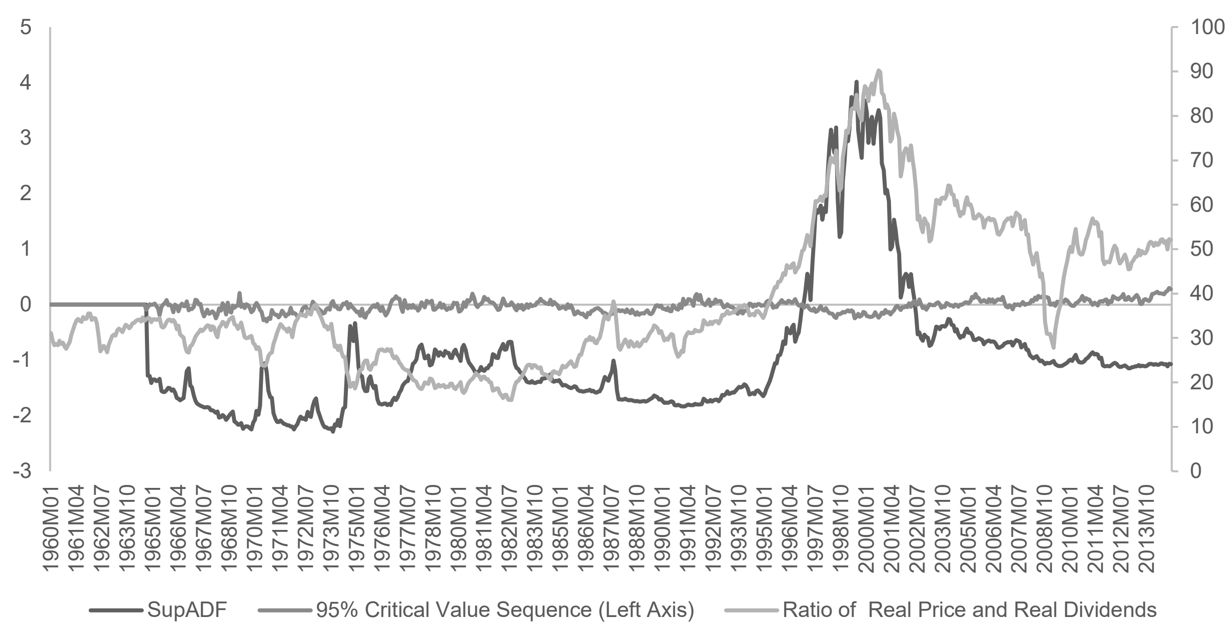
Table 6: The SADF and GSADF Test of the S&P 500 Real Price Dividend Ratio (Sample: Monthly Data 1960:01 - 2014:12)

Test Statistic		**99% level**	**95% level**	**90% level**
SADF	4.01***	2.19	1.73	1.29
GSADF	4.20***	2.74	2.29	2.06

*Critical Values for both tests for three levels of significance are derived from Monte Carlo Simulations with 1,000 replications. The Minimum window size is 53 observations. ***significant at 1 percent level*

The SADF statistic of 4.01 exceeds the 1 percent right tailed critical value (2.19) indicating that the real S&P 500 price dividend ratio experienced explosive periods. To identify a specific bubble period, we compared the backward SADF statistic sequence with 95% critical value sequence obtained from Monte Carlo simulations with 1,000 replications. The results of this are plotted in Figure 2. We can see that SADF test identifies only one bubble, the dot-com bubble. The stock price bubble originated in December of 1997 and terminated in May of 2002.

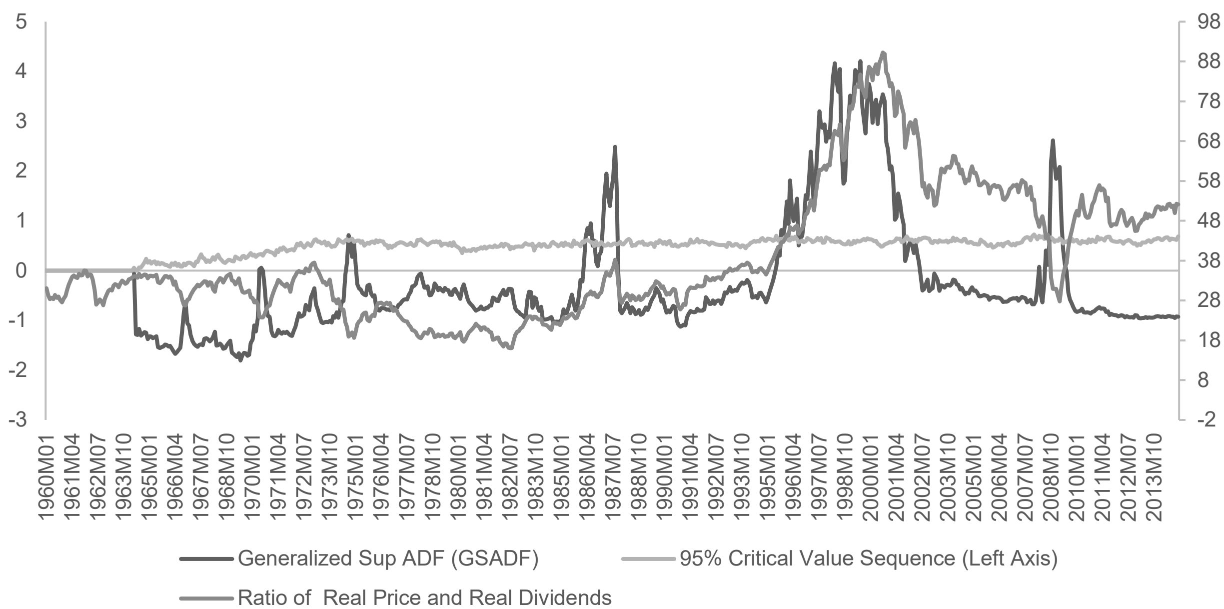
Figure 2: Date Stamping Bubble Periods in the S & P 500 Price-Dividend Ratio: The SADF Test
Sample Monthly Data: January 1960 - December 2014



This figure records the monthly price-dividend ratio for the S&P 500, the values of the SADF statistics and the critical values for the period January 1960 through December 2014. Whenever the SADF value crosses the critical value from below this indicates the start of a bubble. Whenever it crosses from above this signifies the end of a bubble.

Table 6 also presents the results of the GSADF test. This test covers the period 1960-2014. The computational requirements of the GSADF test necessitated testing a shorter period (53 monthly observations) than used by Phillips et al. (2015). The GSADF statistic and the corresponding critical values are presented in Table 6. The GSADF statistic of 4.20 far exceeds the 1% right tailed critical value of 2.74 revealing the presence of multiple bubbles in the sample period. To identify specific bubble periods, BSADF statistic sequence was compared to the 95% SADF critical sequence generated by Monte Carlo simulations with 1,000 replications. The plotted results are shown in Figure 3. The lower panel shows the bubble dating procedure of the GSADF test, the middle line shows the critical value sequence (left axis) and the backward SADF while the right axes measures the S&P 500 price to dividend ratio.

Figure 3: Date Stamping Bubble Periods in the S & P 500 Price-Dividend Ratio: The GSADF Test Sample Monthly Data: January 1960 - December 2014



This figure records the monthly price-dividend ratio for the S&P 500, the values of the GSADF statistics and the critical values for the period January 1960 through December 2014. Whenever the GSADF value crosses the critical value from below this indicates the start of a bubble. Whenever it crosses from above this signifies the end of a bubble.

The Generalized SADF identifies four bubbles between 1960 and 2014. The bubble in 1974, the 1987 black Monday bubble; the dot-com bubble and the subprime mortgage bubble. It is interesting to note that the 1974 bubble is short lived. It started in April of 1974 and ended in September of 1974. The longest identified bubble period is the dot-com bubble. The bubble started in October of 1996 and ended in November of 2001. The black Monday bubble started in November of 1986 and ended in May of 1988. Likewise, the subprime mortgage bubble started in April of 2008 and ended in May of 2009. Interestingly despite the speculation that the recent stock market surge represents a bubble, the GSADF test suggests differently. The backwards SADF statistic values since the end of the subprime mortgage crisis are well below the critical value.

CONCLUDING COMMENTS

This paper's objective is to evaluate common econometric methods available to test for asset price bubbles. We detailed the progress of these tests which first simply tried to detect financial bubbles. To the current state wherein multiple bubbles are detectable and time datable. Therefore, availability of such real time monitoring tools would significantly help investors, retirees, and portfolio managers to rebalance their portfolios during such bubble periods. Similarly, the regulators and the policy makers could adopt appropriate policies to limit the damage to the real economy. We used historical S&P 500 index prices and dividends to employ four of the common methods used to test the presence of bubbles. The variance bound test was one of the first methods used to detect financial bubbles. The test was found to have several problems. This led the development of the West Test. As shown in the paper it was capable of detecting financial bubbles. It could suggest the presence of bubbles but it cannot identify the beginning and ending of a bubble. Cointegration tests examine the time series properties of the data for bubbles. It suffers from the same deficiencies as the West Test that it may suggest the presence of a bubble but cannot identify the starting and ending bubble dates. Phillips et al. (2011) developed right-side unit root tests that are capable of discovering dates of asset price bubbles. However, this test does not identify multiple bubbles and hence Phillips et al. (2015) generalized their initial work by developing a procedure that would identify multiple periodically collapsing bubbles. For example, their generalized procedure identified the formation of a bubble in 1974 and before the 1987 crash. It dated the internet bubble and finally identified the real estate bubble from 2007-2010. While these are generally in line with the perception of the financial press, unlike many in the financial press find no sign of a bubble in the current stock market.

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UNEMPLOYMENT AS A DETERMINANT OF GOLD PRICES: EMPIRICAL EVIDENCE

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ABSTRACT

This objective of this econometrics study is to utilize Pesaran's (2001) Bounds Test to cointegration to study the relationship between gold prices and the US unemployment rate, using three different periods of monthly observation, Model I: 1978-2016, Model II: 1990-2016, and Model III: 2008-2016. Results reveal that there is a long run relationship between the price of gold and unemployment in Models II and III, with Model III representing the strongest and most significant relationship. During 2008-2016 the price of gold increases by 4.7% for every 1% change in the unemployment rate, ceteris paribus. The short run adjustment process however, is stronger between 1990-2016 than between 2008-2016. On the other hand, there is no observed no long run cointegrated relationship between the price of gold and unemployment during the 1978-2016 period. This direct relationship between gold price and unemployment has not been studied in the literature, so further work in this area may lead to greater insight into the impact of this macroeconomic variable on the price of gold.

JEL: E42, L7, N5

KEYWORDS: Cointegration, Gold, Unemployment, Exchange Rate, VIX

INTRODUCTION

Early in the 1800's gold was declared a legal tender. One gold coin equaled $10, and mostly used for large purchases abroad. In the 1900s gold prices were fixed to $35 an ounce under the Bretton Woods agreement. However, with the dissolution of the Bretton Woods agreement in 1973, gold lost its power as a medium of exchange and was no longer required to back up the value of money. Since then, even though central banks continue to hold gold as reserves, it is considered a precious metal whose value is determined by demand and supply forces. The demand for gold arises from consumers, the private sector and government domestically and internationally and the rise for it has been subject to economic cycles, but with an upward trend. The supply of gold however, has stayed relatively flat over time. Since gold is a fixed asset, past gold inventories are still in circulation and increasing in value over time. For example, the cumulative stock of gold in 1974 stood at around 98,000 tons but grew to 175,000 tons by 2015, reflecting roughly a 1.9% growth over the period (World Gold Council, 2016).

Gold Miners and central banks are the most prominent suppliers of gold, with the latter leasing their gold reserves for interest rate since the 1980s (O'Callaghan 1991). Miner's will borrow gold from central banks and sell it on the open market on the assumption that their future extractions would allow them to repay the gold to central banks (Elwell, 2011). There are three major categories for research in gold: its economic and financial prospects, its role as a currency, and the nature and impact of gold mining on the environment and on society (Lucey, 2011; World Gold Council, 2016). From an economic perspective, research has focused on the short-run and long-run determinants of gold prices. Studies have found evidence that U.S Inflation, world inflation volatility, U.S-world exchange rate index, the beta of gold and credit risk default premium all significantly impacted gold prices (Baur & McDermott, 2010). No study however, attempts to analyze the impact of unemployment on the price of gold. This research aims to close this loophole by investigating the impact of unemployment on the price of gold. We proceed in the next section with a review of the literature on gold prices and its determinants. Thereafter we specify our models and variables and explain the data used in estimating our models. The penultimate section explains and discuss our empirical results, while the final section concludes the paper with suggestions for future studies.

LITERATURE REVIEW

One of the most renowned characteristics of gold is the illusion of safety in volatile markets. Baur and McDermot (2009) define gold as a safety asset in short term market turmoil contend that the role of gold in a developed financial system is that of a safe haven asset, but only in the short-run, and only in the presence of extreme volatility. Their study suggests that investors react to short-term, day-to-day volatility, thereby seeking out the safe haven of gold in the short run. An analysis of weekly and monthly stock market losses does not produce the same safe haven refuge response from investors. A study by Mulyadi and Anwar (2012) supports the conclusion of Baur and McDermont (2009) that gold performs as a safe haven asset in short term market turmoil. Mulyadi and Anwar (2012) compare the returns of a stock investment in Indonesian financial markets and an investment in gold. They conclude that when one gets a depreciating return in a stock investment, the return in gold increases. Their results support gold as a diversification instrument and as a safe haven for investors to hedge against stock market risk.

Gold functions as a store of value against inflation. Anssi (2007) examined the short-run and long run determinants of gold prices using Johansen's (1988, 1991) cointegration method. Anssi (2007) found evidence that U.S Inflation, world inflation volatility, U.S-world exchange rate index, beta of gold and credit risk default premium were all statistically significant variables that impact gold prices. A study by Faugere and Erlach (2006) further supports this theory. The authors construct an asset valuation model for gold, dependent on required yield theory, introducing a new exchange rate parity rule. Their model specified that gold prices vary with per capita GDP growth. Real gold prices respond to changes in foreign exchange rates. When the domestic required yield is constant and when foreign exchange rates are constant and there are no major catastrophes, real domestic gold prices increases with domestic inflation. When these factors hold then the real domestic price of gold is determined by the domestic required yield and is impacted by inflation as well as the exchange rate at home. Faugere and Erlach (2006) clearly demonstrate that when exchange rates depreciate gold prices increase, holding everything else constant. In more recent work, Baur (2011) reviewed the relationship of gold with major economic and financial variables and examined whether gold serves as a store of value, influenced by commodity prices, consumer prices, and the value of the U.S. dollar, stock market returns, stock market uncertainty and short and long-term interest rates. The study concludes that gold has evolved as a hedge against financial losses rather than a hedge on inflation. Inferring that when financial instruments deliver depreciating returns, gold takes on a negative correlation allowing one to maintain one's financial wealth.

Qadan & Yagil (2012) observe an interesting connection between gold and the VIX. Their study concludes that variation in investors' sentiment, measured by the VIX, triggers change in the price of gold between January 1995 and May 2010. Their conclusion is significant and opens the door for physiological exams studying the relationship between the mind of investors and how gold prices react. Aggarwal & Lucey (2007) follow this logic and find strong evidence of volatility to returns in the presence of psychological price barriers in gold markets. They conclude that the presence or absence of barriers in gold returns could be a reflection of investor's reactions to changes in interest, inflation and currency markets.

A regression analysis by Fei & Adibe (2010) find a statistically significant relationship between the price movement of gold, real interest rates and the exchange rate, suggesting a close relationship between gold prices and the value of the U.S. dollar. A multiple linear regression study verifies the findings as statistically significant. Furthermore, a study by Başari & Bayramoğlu (2011) yield a negative and statistically significant relationship between the return of gold and the return of the U.S. Dollar. The study also find evidence of a high negative correlation between gold prices and U.S. exchange rates, and a positive correlation between gold prices and oil prices. While GDP has been theorized to have an inverse relationship with gold, empirical studies have yielded mixed results. Lawrence (2003) discovered that gold appeared to be independent of regular business cycles in contrast to other commodities and GDP was uncorrelated with the real rate of return of gold.

On the other hand, an article in the WSJ (December 21, 2012) by Cui & Day, suggested that a drop in gold prices was triggered by an upwards revision by the Commerce Department's of U.S. gross domestic product for the third quarter to 3.1% from 2.7% on December 2012. To date no study has attempted to study the relationship between unemployment and Gold. Unemployment and GDP are related, but it is conventional wisdom that the former is a more commonly accepted indicator of the health of the economy. This was especially true of the financial crises, during which time, unemployment kept rising even though GDP was rising. (The Heritage Foundation, 2010) Unemployment is also a strong indicator of consumer confidence and is often used by politicians to make their case for or against a candidate or political party. Given this, it is important to study the effects of unemployment on gold prices, which is the objective of this paper.

Method, Model Specification, and Data Sources

In order to estimate the price of gold over time, we first specify a general double-log Model I for the period 1978-2016 in Equation (1).

$$lnPG_t = \beta_0 + \beta_2 lnCPI_t + \beta_3 lnSP_t + \beta_4 lnDollar_t + \beta_5 lnUn_t + \varepsilon_t \tag{1}$$

Since 1990, the Chicago Board Options Exchange constructed the VIX, also referred to as the stock market fear index, to quantify stock market volatility over a 30-day period. Qadan & Yagil (2012) observe that variation in investors' sentiment, measured by the VIX, triggers change in the price of gold. However, no further study has been done on this relationship, so this study will integrate VIX into a second general double-log Models II and III for the periods 1990-2016 and 2008-2016 as in Equation (2).

$$lnPG_t = \Psi_0 + \Psi_1 lnVIX_t + \Psi_2 lnCPI_t + \Psi_3 lnSP_t + \Psi_4 lnDollar_t + \Psi_5 lnUn_t + \varepsilon_t \tag{2}$$

In Equations (1) and (2) , PG_t is defined as the monthly London Pm Fix price of an ounce of Gold obtained from the Gold Council and is measured in U.S dollars. The value employed is nominal prices following findings by several studies (Ghazali, Lean, & Bahari, 2015; Naidoo, & Peerbhai, 2015; Ghosh, Levin and Wright, 2004) that the price of gold moves with inflation. SP_t, is the S&P 500, a market value weighted index that measures 500 of the most actively traded stocks in the US financial markets. Empirical evidence of the relationship between PG_t and SP_t is mixed, so we hypothesize that the coefficient for β_2 and Ψ_2 will be indeterminate. $Dollar_t$, is defined as the Nominal Major Currencies Dollar Index with base year 1973. It captures the value of the U.S dollar against major trading partners.

Data are found on the Federal Reserve's database. Empirical evidence of the relationship between PG_t and $Dollar_t$ is mixed (Zheng; Wang & Zheng, 2016), so we expect the coefficient for β_3 and Ψ_3 will be indeterminate. Un_t, The unemployment levels in the economy are measured by the monthly unemployment rate from the Bureau of Labor Statistics. Since no studies have been done to capture the relationship between unemployment and the price of gold, we have no a priori expectations of the sign of its coefficient, so β_5 and Ψ_5 will be considered indeterminate. VIX_t is defined as the volatility index from the Chicago Board Options Exchange. It quantifies the market's expectation of 30-day volatility and is constructed using the implied volatility of a wide range of S&P 500 options. The *VIX* is a widely used measure of market risk and is often referred to as the 'investor fear gauge'. Commensurate with findings in other studies, the relationship between PG_t and VIX_t is negative, so we expect Ψ_1 in Equation (2) to be negative.

We employ the Bounds test model developed by Pesaran, Shin, and Smith (2001) for our cointegration analysis because of its advantages over other models: It does not assume stationarity, or constant means and/or variances of data; it may be applied whether determinants in the model are purely I(0), I(1), or mutually cointegrated; it is able to test the existence of a level relationship between two variables without need to first determine the order of integration of the underlying variables; and it is robust for small and finite samples (Pesaran, *et al.*, 2001; Hendry and Juselius, 2000). The basis of cointegration is that while the response variable and its determinants may be individually non-stationary, they will 'walk' (Murray, 1994) together over time in a cointegrated manner (Engle and Granger, 1987). Furthermore, as other studies have shown, among them, Fama (1965), employing a co-integration model in this study is logical as gold prices, interest rates, and stock markets theoretically follow a random walk and in turn are non-stationary. (Fama, 1965). In estimating the long-run model outlined by Equations (1) and (2), the model will distinguish the short-run effects from the model's long-run dynamics. For this purpose, Equation (1) must be specified in an error-correction model (ECM) format following Pesaran, et al. (2001). Using the Bounds testing approach to cointegration analysis, we rewrite Equation (1) in an ECM format in Equation (3) below.

$$\begin{aligned}\Delta lnPG_t = \alpha_0 &+ \sum_{i=1}^{n} \delta_i \Delta lnPG + \sum_{i=0}^{n} \vartheta lnCPI_{t-i} + \sum_{i=0}^{n} \pi lnSP_{t-i} + \sum_{i=0}^{n} \tau lnDollar_{t-i} \\ &+ \sum_{i=0}^{n} \sigma lnUn_{t-i} + \sum_{i=0}^{n} \varphi lnVIX + \alpha_1 D_{1t} + \lambda_1 lnPG_{i-1} + \lambda_2 lnCPI_{t-1} \\ &+ \lambda_3 lnSP_{t-1} + \lambda_4 lnDollar_{t-1} + \lambda_5 lnUn_{t-1} + \lambda_6 lnVIX_{t-1} + \omega_t\end{aligned} \tag{3}$$

We process two steps in Equation (3) for both Models II and III. The first step utilizes the Wald test to determine the joint significance of the no-cointegration hypothesis H_0: $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5$ (= λ_6 in Model III)=0 against an alternative hypothesis of cointegration H_a: $\lambda_1 \neq 0$, $\lambda_2 \neq 0$, $\lambda_3 \neq 0$, $\lambda_4 \neq 0$, $\lambda_4 \neq 0$ (and $\lambda_6 \neq 0$ in Model III). If the calculated Wald F-value exceeds the upper critical bound value indicating a cointegrated relationship among the explanatory variables, H_0 is rejected; otherwise, H_0 cannot be rejected. The long-run elasticities are the negative of the ECM coefficient of one lagged determinant, for example, Un_{t-1} divided by the coefficient of the lagged response variable, PG_{t-1} to yield the long-run *Un* elasticity of *PG is* (λ_5 / λ_1). The short-run effects are captured by the coefficients ($\delta, \vartheta, \pi, \tau, \sigma, and\ \varphi$ of the first-differenced variables in Equation (3). To estimate Models I, II, and III, monthly data from January 1978 to June 2016 are used. The data series on gold prices are taken from the Gold Council; data on the S&P 500 and the VIX are taken from the Chicago Board Of Exchange found on http://finance.yahoo.com. Data on the CPI and unemployment rate are taken from the Bureau of Labor and Statistics (http://data.bls.gov

EMPIRICAL RESULTS

Cointegration among Variables

Commensurate with Pesaran et al. (2001), Equation (3) goes through two steps. The first step utilizes the Wald test to determine the joint significance of the no-cointegration hypothesis H_0: $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 0$ against an alternative hypothesis of cointegration H_a: $\lambda_1 \neq 0$, $\lambda_2 \neq 0$, $\lambda_3 \neq 0$, and $\lambda_4 \neq 0$. If the calculated Wald F-value exceeds the upper critical bound value indicating a cointegrated relationship among the explanatory variables, H_0 is rejected; otherwise, H_0 cannot be rejected. As can be seen in Table 1, the calculated Wald F-statistic in Model I is below its critical lower bound values (2.003) at the ten percent level, indicating no cointegration between the price of gold and its determinants for the period 1978-2016, leading us to not reject the H_0. Since the relationship between the predictor and response variables are stationary, we drop Model I from our analysis. For diagnostic purposes however, we run the double-log model in Equation 1, and reveals that the unemployment elasticity of *PG* is significant at the 1% level, indicating a strong relationship between the variables when they are not lagged over time.

Table1: Cointegration Results–Gold Price Function, Models I, II &III

Critical Value Bounds of the F-Statistic: Intercept and No Trend						
	10 Percent Level		**5 Percent Level**		**1 Percent Level**	
k	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
4	2.45	3.52	2.86	4.01	3.74	5.06
5	2.26	3.35	2.62	3.79	3.41	4.68

Calculated F-statistic			
Model I (1978-2016);	k =4:	F_{Gold}(PG| SP, Dollar, CPI, U)	2.003
Model II (1990-2016);	k =5:	F_{Gold}(PG| VIX,SP, Dollar, CPI, U)	4.331*
Model III (2008-2016);	k =5:	F_{Gold}(PG| VIX,SP, Dollar, CPI, U)	32.275**

*Note: This table shows the results of the ARDL bounds test for cointegration for different models. Model 1 ranges from 1978-2016 does not include VIX as a regressor, whereas Model 1I ranging from 1990-2016 does include VIX as a determinant of gold. So does Model III, ranging from 2008-2016. Critical values are taken from Pesaran, Shin, & Smith (2001, Table CI(iii) Case III, p. 300). k = # determinants. *, ** indicates statistical significance at the 5% and 1 % level, respectively.*

Models II and III, do however, indicate that cointegration exists between the price of gold and its determinants. The calculated Wald F-statistic for Model II, F=4.331 is above its critical upper bound values at the five percent level, whereas Model III, F = 32.275 is above its critical upper bound value at the one percent level. We therefore reject H_0 in Models II and III, opening the way to estimate the long-run and short-run elasticities of each predictor variable with *PG* for each model.

Long-Run and Short-Run Elasticities, Models II and III

Having established cointegration in Models II and III, we estimate the long-run partial elasticities of the price of gold with respect to each independent variable (Table 2). In addition, we determine the short-run dynamics in Table 3, using the Error Correction Model (ECM). In ECM the movement of any one determinant in time *t,* is related to the gap in time *t-1* from its long-run equilibrium. This step essentially recognizes that economic forces are in constant flux so that the market for gold is seldom in equilibrium. After a shock to the system, ECM facilitates the adjustment back to long-run equilibrium.

Table 2: Gold Price Long-Run Elasticities, Model II: 1990-2016 and Model III: 2008-2016

Independent Variables	**Model II: 1990-2016**		**Model III: 2008-2016**	
	Coefficient	**t-statistic**	**Coefficient**	**t-statistic**
Constant)	-8.439	-2.998***	46.170	3.169***
lVIX	0.141	1.356	-1.798	-6.995***
lSP	0.627	-2.560**	2.241	3.049***
lDollar	-1.354	-2.144**	6.743	8.259***
lCPI	4.910	4.514***	-12.937	-3.872***
lU	0.216	0.862	4.688	5.876***
Adjusted R-squared (R^2)	0.33		0.86	

*Note: This table shows the long-run elasticities of the estimated gold price function for the periods 1990-2016 and 2008-2016. Here the dependent variable is the natural log of the price of gold, PG. Independent variables are listed in the table. *** and ** indicate statistical significance at the 1% and 5% level, respectively.*

Table 2 shows the long-run elasticities between the price of gold and its determinants for Model II: 1990-2016, and Model III: 2008-2016. The signs of the coefficients of all independent variables in Model II, VIX, SP, Dollar, CPI, and Unemployment, meet our research expectations. SP and Dollar are significant at the 5% level, whereas CPI, whose partial elasticity to PG is 4.9, is significant at the 1% level. However, VIX, which is positive and inelastic, is not significant, which is supported by recent studies. Of special interest here is the partial elasticity of unemployment with respect to PG. It is positive and inelastic with a 10% increase leading to an increase in PG by 2.16% , holding other variables in the model constant. However, this occurs at the 30% significance level only, which renders weak results. The S&P 500 is significant and yields a positive sign, signifying that an increase in the value of the stock market by 10% should increase the price of gold by 6.27%, *ceteris paribus*. Adjusted R^2 is 0.33.

Model III, unlike Model II, yields excellent results. All independent variables are significant at the 1% level. We note in particular, that since the financial crisis period is heavily represented in Model III, the coefficients of all the determinants of PG are highly elastic. In the case of unemployment in particular, its partial elasticity with respect to PG is positive and highly elastic with a 1% change in Un positively affecting PG by 4.87%, *ceteris paribus*. Similarly, in the case of VIX, an increase of 1% leads to a decrease in PG of 1.8% indicating that volatility in the stock market leads to purchasing more gold and therefore, driving up its prices, *ceteris paribus*. Adjusted R^2 is 0.86, indicating the PG is very well explained by its determinants.

We present the estimated short-run elasticities between the price of gold and its determinants for Model II: 1990-2016, and Model III: 2008-2016 in Table 3. Results suggest that the short run partial elasticities of all the determinants with respect to their response variable, PG, are statistically significant, mostly at the 1% level for both Models II and III. Model II produces a negative sign for the VIX indicating that a decrease of 1% in the change in VIX leads to an increase to a change in PG of 0.03%, and it is significant at the 1% level. In contrast, in Model III, which encompasses the 2008 financial crisis, an increase of 1% in the change in VIX leads to an increase to the change in PG of 0.17%, and it is significant at the 1% level. Since the VIX measures volatility in financial markets, but does not differentiate between positive and negative volatility, a lot of volume in financial markets can lead to higher or lower stock prices; hence our results can be expected. Our variable of interest, Unemployment, in both models are statistically

Table 3: P Error-Correction Model and Short-Run Elasticities, Model II: 1990-2016 and Model III: 2008-2016

Model II: 1990-2016			**Model III 2008-2016**		
Independent Variable	**Coefficient**	**t-statistic**	**Independent Variable**	**Coefficient**	**t-statistic**
Constant	0.000	0.000	Constant	0.000	0.000
$\Delta \ln PG_{t-11}$	0.144	2.873***	$\Delta \ln PG_{t-9}$	0.165	2.803**
$\Delta \ln VIX_{t-12}$	-0.027	2.048**	$\Delta \ln VIX_{t-6}$	0.367	13.493***
$\Delta \ln SP_{t-2}$	0.102	2.363**	$\Delta \ln SP_{t-7}$	0.288	3.148***
$\Delta \ln Dollar_{t}$	-1.168	-7.124***	$\Delta \ln Dollar_{t-2}$	-0.790	-2.678**
$\Delta \ln CPI_{t}$	2.545	3.568***	$\Delta \ln CP_{t-12}$	5.412	4.298***
$\Delta \ln U_{t-11}$	-0.179	-2.623***	$\Delta \ln U_{t-6}$	0.426	3.509***
ecm1	-0.061	-5.148***	ecm1	-0.201	-15.730***
Adjusted R-squared (R^2)			Adjusted R-squared (R^2)		

*Note: This table shows the results of the short-run partial elasticities of the estimated gold price function for the periods 1990-2016 and 2008-2016. Here the dependent variable is the change in the natural log of the price of gold, PG. *** and ** indicate statistical significance at the 1% and 5% level, respectively.*

significant at the 1% level. In Model II, the $\Delta \ln U_{t-11}$ takes on a negative relationship with $\Delta \ln PG$, with a 10% decrease in $\Delta \ln U_{t-11}$ leading to a 1.7% increase in $\Delta \ln PG$. However, Model III capturing their relationship during the financial crisis, supports our hypothesis, that unemployment is a determinant of PG. The error correction term, ECM_{t-1}, gauges the rate at which PG adjusts to short-run deviations in its determinants, VIX, PG, U, SP and Dollar, before returning to its long run equilibrium level. The coefficient for ECM_{t-1} in both models is negative as is theoretically expected, signifying that the system is converging to equilibrium. Values of -0.061 in Model II and -0.201 in Model III, indicate that once the model in Equation (2) is shocked, convergence to equilibrium is 6% and 20% respectively, of the adjustment occurring within the first month. Clearly, Model III has a far more rapid response rate than Model II, although both models are adjusting quite slowly. Although it is assumed that the ECM complies with the classical normal linear regression model specifications, including no serial correlation and no perfect multi-colinearity, and that the model is correctly specified, we perform the following diagnostic tests, namely, the Durbin-Watson, Breusch-Godfrey (Basak, et al., 2012), RESET and the Augmented Dickey-Fuller, to test these hypotheses. Results for both Model II and Model III are shown in Table 4.

Table 4: Results of the Diagnostic Tests

Diagnostics	**Model II: 1990-2016**		**Model III: 2008-2016**	
	Coefficient	**p-value**	**Coefficient**	**p-value**
R-squared	0.38	--	0.97	--
Adjusted R-squared	0.34	--	0.91	--
Durbin Watson Test	1.902	0.154	2.514	0.910
Breusch-Godfrey Test	0.6	0.663	0.983	0.440
RESET Test	3.6069	0.03	0.004	0.99
Augmented Dickey-Fuller	-2.153	0.51	-2.4138	0.405

Note: This table shows the diagnostic tests for Models II and III to determine the presence of serial correlation, multi-colinearity, and correct model specification.

In Table 4, the Durbin-Watson (D-W) test is used to test autocorrelation in the residuals, yields a p-value of 0.154 in Model II and 0.910 in Model III, so we do not reject the null hypothesis of no autocorrelation in both models. The Breusch-Godfrey test is appropriate in the presence of stochastic regressors such as lagged values of the dependent variable for higher order autocorrelation, and is asymptotically equivalent to the Durbin-Watson test for first order autocorrelation (Rois, Basak, Rahman, Majumder 2012). The test results in a p-value of 0.663 in Model II and 0.440 in Model III, reinforcing the D-W test results of no serial correlation in the errors of the model. The RESET test, used to determine if the model is correctly specified, produces non-significant values in Model III,, leading to a non-rejection of the null hypothesis. The Augmented Dickey-Fuller Test in both models yield non-significant results leading to a non-rejection of the null Hypothesis of the absence of a unit root; that is, both models are non-stationary.

Conclusions, Limitations, and Suggestions for Future Research

This study attempted to model gold prices as a function of the unemployment rate using Pesaran's (2001) cointegration model. Three models were created under different conditions: Model I encompassed the period 1978-2016, in which the price of gold was regressed against unemployment, the dollar exchange rate, the Consumer Price Index, and the S&P 500. Model II covering the period 1990-2016 included the VIX as a determinant of the price of gold. Model III replicated Model II but for the period 2008-2016 to capture the effects of the global financial crisis of 2008-2010. Results from Models II and III indicate that cointegration exists between the price of gold and its determinants. This is not so for Model I, leading to us dropping the model. In the long run, in both models, the estimates of the partial elasticity of unemployment is positive; however, Model II renders non-significant results, whereas Model III indicates a strong positive and elastic relationship between unemployment and the price of gold. Short-run elasticities in Model II are negative, inelastic, and significant at the 1% level, whereas in Model III it is positive, inelastic, and significant at the 1% level. It is Clear that the 2008 financial crisis, strongly represented in Model III, reflects a strong relationship between unemployment and the price of gold.

Diagnostic test results indicate that models are neither serially correlated, nor stationary. Model III is correctly specified and displays correct functional form. Adjusted R^2 in Model III in the short and long runs indicate that variation in the price of gold is explained by its determinants. This is, however, not the case for Model II, implying that that the 2008-2016 Model is a better indicator of the relationship between unemployment and the price of gold.

To our knowledge, this is the only study to date that estimates the link between the price of gold and the unemployment rate in the US. As indicated above, our results are mixed, with Model III showing robustness. Future studies may revisit the unemployment-price of gold relationship by using quarterly rather than monthly data and by including real GDP as a determinant of the price of gold. Since the price of gold has been linked to consumer confidence, it may be helpful for future studies to incorporate the Consumer Confidence Index, which quantifies the degree of consumer optimism in the economy, into the models. Furthermore, Model I could include a dummy variable to capture the various business cycle downturns and troughs between 1978 and 2016.

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FORECASTING VOLUME AND PRICE IMPACT OF EARNINGS SURPRISES USING GOOGLE INSIGHTS

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ABSTRACT

This paper examines the predictability of price and volume movements using Google Insights on equities exhibiting earnings surprise and the association with pre-announcement information searching. The motivation for this paper is to answer two primary research questions. First of all, using more recent stocks earnings surprise, is Google search data a good indicator of investor interest prior to the earnings announcement? Second does the Google data add to the predictability of post earnings volume and pricing direction? Data on earnings surprise were taken from Yahoo Finance and Google search volume data were taken from the Google trends website. While the results found in the analyses above are not highly convincing regarding Google trends data and price movement from earnings surprise, the results on the volume models yielded promising (i.e. significant) results. Moreover, Mean Absolute Error was reduced by approximately 8% when incorporating the Google trends data on volume predictions.

JEL: G17

KEYWORDS: Predictability, Volume Movements, Earnings Surprise, Google

INTRODUCTION

Financial research suggests that the impact of earnings surprise on firms quoted price is a function of investor's expectation of the surprise, as well as their reaction to the surprise itself (c.f. Atiase and Bamber, 1994 and Barberis, Shleifer and Vishny, 1998). This implies that pre-earnings expectations will absorb some of the price movements associated with the actual earnings surprise. An unanticipated announcement on the other hand can have both immediate and persistent effects on the market clearing price. The ability to predict (1) the earnings surprise and (2) the price and volume impact on the particular asset has been the focus of considerable research (see for example Barberis, Shleifer and Vishny, 1998). Among the many studies examining this phenomenon researchers have employed event studies, single factor and multi factor models. (c.f., Fama (1991) and Fama and French (1996), Mackinlay (1997)). This paper examines the predictability of price and volume movements using Google Insights on equities exhibiting earnings surprise and the association with pre-announcement information searching.

LITERATURE REVIEW

Considerable work has been done on stock earnings surprise and abnormal returns. For example, Fama (1998) found that finance literature identified many long-term return anomalies that were inconsistent the efficient market hypothesis (EMH). These anomalies are found to exist both pre and post earnings announcement, as well as during merger activity (Wansley, Lane and Yang, 1983). As financial markets have grown, both domestically and internationally, investors and researchers have continued to focus on identifying anomalies in order to earn excess short-term returns (DellaVigna and Pollet, 2009). This particular line of inquiry has influenced (and been influenced by) research that seeks to identify and measure information demand of investors around earnings announcements. Vlastakis and Markellos (2012) studied volatility and information demand. They found that information demand at the market level is positively related to historical and implied measures of volatility. Furthermore they find that information demand increases during periods of higher returns. This is consistent with the idea that trading volume and subsequent price movements react to quarterly earnings announcements (Bamber, 1987).

Drake, Roulstone and Thornock (2012) examined information demand of investors around earnings surprise using Google search index data. They find that the information build up around the event begins approximately two weeks prior to the earnings announcement and continues beyond the earnings announcement. They also find that part of the earnings surprise is already incorporated into the price of the stock prior to earnings; therefore, the price impact of the surprise is diffused around the announcement. The specific motivation in this paper was to assess both the nature and timing of investor information demand (and the pricing impacts) during earnings season. The authors employ a number of regression models that attempt to predict search volume and abnormal returns during varying time windows. The concept of information demand and investor information seeking has been of interest over the last 40 years. As information becomes more accessible, additional data sources available to the public have begun being used. These sources include Facebook, Twitter and Google.

Google trends shows the search volume of a particular topic over a particular point in time. Google describes Google Trends on their website, and define the numbers on the graph as reflecting how many searches have been done for a particular term, relative to the total number of searches done on Google over time. They don't represent absolute search volume numbers, because the data is normalized and presented on a scale from 0-100. Each point on the graph is divided by the highest point, or 100. When we don't have enough data, 0 is shown (https://support.google.com/trends/answer/4365533?hl=en)." Choi and Varian (2009) conducted a broad series of analyses in the paper "Predicting the Present with Google Trends." In this paper the authors employ a seasonal autoregressive model to predict automobile sales, home sales and travel. In doing so they found that using Google search data improved predictive power over an autoregressive model with a single lag parameter. Other research using these data have also been developed given the unique characteristics of Google's data. For example, additional research has been conducted in finance and economics, as well as, epidemiological studies (c.f., Pelat et. al. (2009), Preis et. al. (2013)).

The availability and accessibility of Google search data to researchers has provided an interesting and innovative direction for different types of research. With respect to financial research, and information seeking specifically, Google search data can be seen interpreted as information seeking by individuals and, in the case of this particular analysis, investors. Bushee, Core, Guay and Hamm (2010) recently addressed this topic whereby they examine the impact of the business press on reducing information asymmetry. Their findings indicate that when the media provides information about the potential for surprise, it reduces the price and volume impact for a particular asset. This is a similar result that was found by Drake et al (2012) where they concluded that the Google trends data is a form of information seeking by investors, which reduces pre-announcement information asymmetry.

The motivation for this paper is then to combine the focuses of the three aforementioned papers. There will be two primary research questions. First of all, using more recent stocks earnings surprise, is Google a good indicator of investor interest prior to the earnings announcement? Second does the Google data add to the predictability of post earnings volume and pricing direction? If robust results are found for these questions then the question regarding information asymmetry reduction will provide validation for the Bushee et al (2010) paper.

DATA AND METHODOLOGY

Data on earnings surprises were taken from yahoo finance (http://biz.yahoo.com/z/extreme.html). Here you can search for earnings dates and whether or not there was an earnings surprise associated with a particular equity. In this paper companies were selected over a three day period in August 2012. This period was selected in order to capture a one year time horizon in which the companies analyzed did not have a prior earnings surprise in the prior one year. This article was initially written in September, 2013 and the earnings season from August 2013 was the most relevant. For each day five equities were selected; two with positive earnings surprises (earnings per share (EPS) above analysts' consensus estimates), two with negative earnings surprises (earnings per share (EPS) below analysts' consensus estimates and one that met earnings expectations (i.e. no surprise). Assets with earnings dates in August 2013 were selected. They were further segmented into those with positive, negative and no earnings surprise in order to select from among those groups. Thereafter, assets were selected randomly from within each earnings direction group, irrespective of industry, company size, market capitalization, etc. Furthermore, firm performance (i.e. book to market, P/E ratio, cash flow to price, etc.) was not considered in this study, although the direction and magnitude of surprise can be indicative of current firm performance (Drake et al, 2012). Finally, weekly data were captured for the preceding 52 weeks closing prices and volumes leading up to the earnings announcement in which the surprise occurred. The only caveat to the selection of the assets was that they must not have had any other earnings surprises in the prior 12-month period.

Google interest data were taken from the Google trends website (http://www.google.com/trends/). Google trends data allows users to search a particular keyword over a particular time interval. The interest data are then standardized in order to index relative search volume (not absolute search volume) over time. Term searches are unstructured in the Google trends environment. One can search any term and identify whether or not search volume occurred over the specified time interval. Similarly, when searching for company information, entering the entire company name will generate interest data that are not exclusively search entries that were information gathering as a result of earnings expectations. In order to control for this potential contemporaneous interest result, search terms were only entered as ticker symbols (c.f. Drake, Roulstone and Thornock (2012)). Table 1 shows the companies that were used in the analysis, ticker symbols, surprise type, surprise magnitude, analyst consensus expectation for EPS and actual reported EPS.

Table 1: Companies Used in Analysis and Summary Statistics of Earnings Surprise

Company Name	Symbol	Surprise Date	Earnings Surprise Direction	Surprise %	Reported EPS	Consensus EPS
Medtronic	MDT	8/20/2013	Met	0.0%	0.88	0.88
Best Buy	BBY	8/20/2013	Upside	166.7%	0.32	0.12
Home Depot	HD	8/20/2013	Upside	2.5%	1.24	1.21
Dicks Sporting Goods	DKS	8/20/2013	Downside	-4.1%	0.71	0.74
JC Penney's	JCP	8/20/2013	Downside	-107.6%	-2.2	-1.06
Smuckers	SJM	8/21/2013	Upside	3.3%	1.24	1.20
Lowes	LOW	8/21/2013	Upside	11.4%	0.88	0.79
Hewlett Packard	HP	8/21/2013	Met	0.0%	0.86	0.86
Eaton Vance	EV	8/21/2013	Downside	-3.7%	0.52	0.54
Staples	SPLS	8/21/2013	Downside	-11.1%	0.16	0.18
Prospect	PSEC	8/22/2013	Upside	26.7%	0.38	0.30
Pandora	P	8/22/2013	Upside	100%	0.04	0.02
Gap	GPS	8/22/2013	Met	0.0%	0.64	0.64
Abercrombie	ANF	8/22/2013	Downside	-42.9%	0.16	0.28
Sears Holdings	SHLD	8/22/2013	Downside	-54.6%	-1.70	-1.10

Table 1: Shows companies used in the analysis, ticker symbols, earnings dates and earnings results during August 2013.

In table 1, a surprise type of "Met" indicates EPS expectation was realized in the actual reporting, "Downside" indicates EPS was below expectation and "Upside" indicates EPS expectations were above expectation. Companies were selected that had met expectations as a reference group. This is consistent with some research in the event study literature (c.f., Lee, 2007). Although this is not an event study methodology, the ability to determine a baseline of predictability using companies that had met earnings expectations was utilized as a means of comparison. Table 2 shows the summary statistics for each of the equities used in the analysis, including their surprise, mean, median and range of closing prices, traded volume and Google interest.

METHODOLOGY

The methodology being employed is similar to that in the Choi and Varian (2009) paper. In this paper Choi and Varian employ a basic autoregressive model to estimate various macroeconomic factors mentioned previously. Their results indicate that the when adding Google trends data to the autoregressive model, overall prediction error is reduced significantly.

Table 2: Company, Surprise Direction and Weekly Price, Volume and Google Interest

Company (Ticker)	Surprise	Observations	Mean Weekly Close	Mean Weekly Volume	Mean Weekly Interest
ANF	Negative	52	$45.23	2,358,440	79
BBY	Positive	52	$20.80	8,553,342	60
DKS	Negative	52	$49.81	1,418,013	62
EV	Negative	52	$35.96	866,242	73
GPS	No Surprise	52	$37.15	4,594,827	48
HD	Positive	52	$68.96	7,381,806	53
HPQ	No Surprise	52	$19.43	24,451,706	39
JCP	Negative	52	$19.00	12,752,879	34
LOW	Positive	52	$37.63	9,604,185	81
P	Positive	52	$13.01	5,960,602	56
PSEC	Positive	52	$11.05	3,058,562	42
SHLD	Negative	52	$49.15	1,008,440	48
SJM	Positive	52	$95.25	594,335	64
SPLS	Negative	52	$13.36	11,143,367	44

Table 2: Shows the direction of the earnings surprise, the number of observations evaluated and the weekly mean closing price, volume traded and Google interest.

A similar autoregressive model will be built in this analysis as well, using both prior week's closing prices and volume traded. The model will be used, to forecast pricing and volume, before and after Google trends data are entered as predictors in the model. Secondarily, it will examine the impact of the surprise on the direction of the volume and price movements. For example, pre and post earnings movement exists has been examined extensively in the literature and is referred to as "earnings announcement drift" (c.f. Bernard and Thomas, 1989). Investors will take positions in assets prior to (and post) earnings announcements in order to capitalize on the added volatility (c.f. Sadka, 2006). In the model we will examine the impact of Google search volume as a predictor of volume and closing price in the following week. The hypothesized results will be as follows:

Hypothesis 1-Google trends data will help identify an increase in volume traded for both positive and negative surprise stocks.

Hypothesis 2-Google trends data will help predict the direction of the price change for positive and negative surprise shocks.

*Overall prediction error in both volume and closing price will be reduced as measured by mean absolute error (MAE) by adding Google trends search data. This expected result is consistent with the Choi and Varian (20*09) outcome.

The model specification for the baseline approach will be,

$$volume_t = \beta_0 + \beta_1(volume_{t-1}) + \epsilon_t \quad (1)$$
$$close_t = \beta_0 + \beta_1(close_{t-1}) + \epsilon_t \quad (2)$$

Where, subscript t and t-1 denote current and prior week cumulative values for volume and ending values for price, respectively ϵ_t represents the error term in both models.

Once the baseline coefficients and errors are calculated, the Google interest data will also be included and the following models will be estimated,

$$volume_t = \beta_0 + \beta_1 \log(volume_{t-1}) + X_i + \epsilon_t \quad (3)$$
$$close_t = \beta_0 + \beta_1(close_{t-1}) + X_i + \epsilon_t \quad (4)$$

Where, variables defined previously are the same and the X_i refers to the Google interest data indexed over the prior year. From here the revised parameter estimates will be examined and MAE will be calculated for the models with Google trends data. The results will be compared to the hypotheses listed above. It is expected that the results for volume should always be positive on the Google coefficient. An expectation of a surprise will lead to more volume traded irrespective of whether or not they are sold or bought. Where surprises (positive or negative) exist the coefficient should be significant. Price expectations on the other hand should be directional. Therefore, a stock with positive earnings surprise should be identified by a positive Google interest coefficient and a negative earnings surprise should be identified by a negative coefficient on the Google interest variable.

RESULTS

The results of the various Google interest data were most revealing on volume predictions for stocks with the highest volume and/or stocks with the largest surprises (irrespective of surprise direction). For example, JCP, SHLD and SPLS had significant volume and close coefficients on the Google interest variable. In these assets, both JCP and SPLS have significantly higher median weekly volume, relative to the average of the volume for the entire sample (12.7MM and 11.1MM shares compared with 6.6MM shares for all). SHLD is an anomaly due to below average volume (1MM shares); it has significant Google interest coefficients on both volume and price, although likely due to large funds eliminating remaining positions and/or small and medium sized traders capitalizing on volatility due to the firm's financial distress. This movement in SHLD is more than likely a result of the considerable negative news that has been reported recently. The downward spiral to bankruptcy has been observed in other firms where financial distress has been long term and with significant magnitude (c.f. Gilbert and Menon, 1990). The remainder of the results are quite fragmented (i.e. significant coefficients on Google interest and closing price for no surprise stocks). Additionally, there are a number of assets with sufficient volume and large enough earnings surprise to generate significance that did not. All coefficients on volume and price can be seen in sections 1 and 2 of the appendix. In addition, plots of the predictions (both with and without Google interest) are compared to the actual volumes and closing prices in sections 3 and 4 of the appendix. Generally, the predictions with interest are much closer to predictions without interest particularly in stocks where the Google trends data were significant on both price and volume. This brings us to our third hypothesis, which is whether or not including Google trends data into the model would reduce the mean absolute error (MAE). The MAE is defined as

$$MAE = \frac{1}{n}\sum_{i=1}^{n} |\hat{y}_i - y_i| \quad (5)$$

Where, $\hat{y}_i$ represents the forecast of y_i and y_i represents the actual result. Tables for the MAE on all assets are listed in Table 1 below. Overall, the MAE is reduced from 29.6% to 27.3% for volume and effectively unchanged for closing prices. In order to assess whether or not there was a statistically significant difference in these two values a paired samples t-test was run and the results are listed in Table 3 below.

Table 3: Comparison of Mean Absolute Error (Weekly Volume and Close) by Company

	Volume Error No Interest	Volume Error w/ Interest	Close Error No Interest	Close Error w/ Interest
ANF	30.8%	32.0%	4.1%	4.0%
BBY	44.8%	43.6%	5.1%	5.1%
DKS	25.6%	27.1%	2.2%	2.1%
EV	28.7%	28.7%	2.9%	3.0%
GPS	27.7%	25.5%	2.5%	2.5%
HD	19.0%	19.1%	2.0%	2.1%
HPQ	34.2%	19.7%	4.0%	4.0%
JCP	34.2%	26.6%	6.1%	6.3%
LOW	22.2%	22.1%	2.2%	2.2%
P	33.5%	30.5%	5.9%	5.8%
PSEC	27.2%	29.1%	1.6%	1.6%
SHLD	33.9%	26.4%	5.1%	5.1%
SJM	19.0%	19.2%	1.4%	1.5%
SPLS	34.2%	31.9%	2.5%	2.5%
MEAN	29.64%	27.25%	3.40%	3.42%
Median	29.75%	26.86%	2.72%	2.76%
Minimum	18.96%	19.07%	1.44%	1.46%
Maximum	44.82%	43.64%	6.09%	6.33%

Table 3-Mean Absolute Error for each firms price and volume forecasts with and without Google interest. In the table above "No Interest" is without including Google data and "w/Interest" is with Google data. The final four rows of data are descriptive statistics for each models error.

Table 4 below shows the results of the t-test for the MAE difference in predictive power between models with and without Google interest included.

Table 4: Paired Samples T-Test's for Mean Absolute Error Before and After Including Google Trends in Model

Statistics	Volume Error No Interest	Volume Error w/ Interest	Close Error No Interest	Close Error w/ Interest
Mean	0.2963	0.2725	0.0342	0.034
Variance	0.0048	0.0042	0.0002	0.0003
Observations	14	14	14	14
Pearson Correlation	0.7741		0.9989	
Hypothesized Mean Difference	0		0	
Df	13		13	
t Stat	1.96		1.288	
P(T<=t) one-tail	0.0358		0.11	
t Critical one-tail	1.77		1.771	
P(T<=t) two-tail	0.0717		0.22	
t Critical two-tail	2.16		2.16	

Table 4-Paired two sample t-test comparing differences in group means between models with and without Google trends data. Statistically significant difference exists in MAE on volume prediction differences (lower MAE with Google trends data than without). Difference in MAE on close prediction differences is not statistically significant.

The mean MAE for closing price indicated no meaningful difference between the models (3.42% to 3.40%). The paired samples t-test found that the difference between these two groups was not statistically significant. The difference on volume on the other hand was found to be statistically significant at the 5% level (p-value .035 and t-statistic 1.96). When re-examining the coefficients (appendix sections 1 and 2) it is clear that 9 of the 14 firms examined had statistically significant parameter estimates on the Google trends data variable in the volume models compared to 4 of 14 for the closing price models.

CONCLUDING COMMENTS

This paper attempted to apply the methodology from Choi and Varian (2009) to stock earnings surprise. Predicting volume and pricing using Google trends data appears to provide some lift in prediction accuracy of forecasts. The focus of this paper was to examine the efficacy of Google insights data as predictors in a model. Because the data are an index of the time period of interest we do not truly see absolute search volume, but only relative minimum and maximum values over our specified time interval. The results found an increased predictive power, via a lower MAE, for the volume forecast. Results on price prediction were not as promising. For example, the asset selection process in this paper was very subjective. Assets were selected based on an arbitrary time period (the most recent earnings season) and were then filtered according to the non-existence of earnings surprise in the prior year. A more robust selection process on a larger number of assets may yield more insightful results. Furthermore, only a one year time horizon was examined using weekly data. The frequency of the data and/or the time horizon may also have been limiting factors in generating significant (and consistent) results. The fact that MAE reduction did occur with statistical significance is promising and should provide a good starting point for examination in future research.

APPENDIX

Appendix 1: Volume Data

Company and Model Used	Intercept	Lag of Volume	Google Interest
ANF (w/o Google Interest)	2,103,453.13***	0.12	
ANF (w/ Google Interest)	-3,195,523.26	0.12	67303.03
BBY (w/o Google Interest)	5,031,934.41***	0.41***	
BBY (w/ Google Interest)	8,826.72	0.4***	85955.1**
DKS (w/o Google Interest)	1,246,790.78***	0.13	
DKS (w/ Google Interest)	-1,174,675.54**	0.12	38968.31***
EV (w/o Google Interest)	512,310.58***	0.41***	
EV (w/ Google Interest)	44,1947.98	0.42***	883.29
GPS (w/o Google Interest)	2,609,322.32***	0.43***	
GPS (w/ Google Interest)	426,065.06	0.41***	47980.93**
HD (w/o Google Interest)	5,870,569.89***	0.21	
HD (w/ Google Interest)	5,344,977.36***	0.18	13601.76
HPQ (w/o Google Interest)	17,743,397.61**	0.28*	
HPQ (w/ Google Interest)	-3,693,140.62	0.08	674538.98***
JCP (w/o Google Interest)	4,430,108***	0.68***	
JCP (w/ Google Interest)	-155,759.68	0.15*	323134.12***
LOW (w/o Google Interest)	558,349.27***	0.42***	
LOW (w/ Google Interest)	464,9445.24	0.41***	11998.26
P (w/o Google Interest)	3,469,315.71***	0.41***	
P (w/ Google Interest)	-3,386,505.07*	0.33***	130810.55***
PSEC (w/o Google Interest)	2,628,680.79***	0.15	
PSEC (w/ Google Interest)	194,408.87**	0.05	64528.53***
SHLD (w/o Google Interest)	651,166.94***	0.35**	
SHLD (w/ Google Interest)	-221,004.64	0.3**	19309.58***
SJM (w/o Google Interest)	365,644.65***	0.38***	
SJM (w/ Google Interest)	316,388.16**	0.38***	742.37
SPLS (w/o Google Interest)	7,872,009***	0.3**	
SPLS (w/ Google Interest)	-2,007,832.66	0.25**	240421.48***

*Appendix 1 shows the parameter estimates for volume using prior week's volume lag and Google interest data. ***, ** and * indicate significance at the 1%, 5% and 10% levels.*

Appendix 2: Price Model

Company and Model Used	Intercept	Lag of Closing Price	Google Interest
ANF (w/o Google Interest)	4.58*	0.9***	
ANF (w/ Google Interest)	1.14	0.89***	0.05
BBY (w/o Google Interest)	-0.2	1.03***	
BBY (w/ Google Interest)	-0.63	1.02***	0.01
DKS (w/o Google Interest)	12.89**	0.74***	
DKS (w/ Google Interest)	13.26**	0.75***	-0.01
EV (w/o Google Interest)	2.47*	0.94***	
EV (w/ Google Interest)	1.99	0.92***	0.02
GPS (w/o Google Interest)	1.07	0.98***	
GPS (w/ Google Interest)	1.53	0.98***	-0.01
HD (w/o Google Interest)	4.21*	0.94***	
HD (w/ Google Interest)	5.71**	0.89***	0.03**
HPQ (w/o Google Interest)	0.67	0.97***	
HPQ (w/ Google Interest)	1.29	0.96***	-0.01
JCP (w/o Google Interest)	1.63	0.9***	
JCP (w/ Google Interest)	5.24***	0.78***	-0.04***
LOW (w/o Google Interest)	1.42	0.97***	
LOW (w/ Google Interest)	2.7	0.97***	-0.02
P (w/o Google Interest)	0.24	0.99***	
P (w/ Google Interest)	-0.15	0.97***	0.01
PSEC (w/o Google Interest)	2.16**	0.8***	
PSEC (w/ Google Interest)	2.28**	0.8***	0
SHLD (w/o Google Interest)	6.33	0.87***	
SHLD (w/ Google Interest)	10.71***	0.88***	-0.1***
SJM (w/o Google Interest)	1.51	0.99***	
SJM (w/ Google Interest)	1.53	0.98***	0.01
SPLS (w/o Google Interest)	0.88	0.94***	
SPLS (w/ Google Interest)	1.39***	0.96***	-0.02***

*Appendix 2 shows the parameter estimates for price using prior week's price lag and Google interest data. ***, ** and * indicate significance at the 1%, 5% and 10% levels.*

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COULD BASEL III CAPITAL AND LIQUIDITY REQUIREMENTS AVOID BANK FAILURE?

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ABSTRACT

The aim of this study is to examine the contribution of the Basel III requirements in reducing bank failure risk through three different measures: the new long-term liquidity ratio (Net Stable Funding Ratio: NSFR), the Leverage ratio and the capital Tier One ratio. We use data on U.S. commercial banks during the 2008-2010 subprime crisis period. Our results depend on bank size: small banks are more sensitive to their fundamentals than large banks when it comes to failure risk. For large banks, no more safety is driven from the Leverage ratio or from the NSFR when Tier One ratio is applied. We also find that Leverage ratio considering off-balance sheet can be a complementary constraint for reducing bank regulatory arbitrage.

JEL: G01, G20, G21, G28

KEYWORDS: Financial Crisis, Bank Failure, Liquidity Creation, Basel III, Regulation, NSFR

INTRODUCTION

Basel III emerged in response to the 2007-2008 subprime crisis. To reduce systemic risk, regulators consider both the asset risk through the capital requirements and the transformation risk through two liquidity ratios (one for the short-term, Liquidity Coverage Ratio, LCR, and one for the long-term, Net Stable Funding Ratio, NSFR). Previous literature on failure risk focused either on asset risk or on transformation risk. The first theory considers asset risk as the main cause of bank failure and uses CAMEL (Capital adequacy, Asset quality, Management, Earnings, Liquidity and Sensitivity to market risk) indicators as a tool of analysis (Lanine and Vander Vennet, 2006; Wheelock and Wilson, 2000). The second theory supports the idea that transformation risk can cause financial instability leading to systemic crisis. In fact, there is a mismatch between liquid short-term payable deposits, and illiquid long-term loans (Diamond and Dybvig (1983). Few studies tried to explore simultaneously the role of these two risks in bank failure (Vazquez and Federico, 2012).

Some of empirical studies explored the impact of Basel III requirements but did not consider the off-balance sheet. Our paper will focus on the major pillars of the Basel III requirements (capital and liquidity) being essential to avoid bank failure and financial crises. It will also bring up the idea that size can affect bank sensitivity to these requirements. This paper will be mainly divided in three major parts. Section 1 will be the literature review. Section 2 will focus on data and the methodology applied. In section 3, we will analyze and evaluate applying Basel III requirements. Finally, we will conclude.

LITERATURE REVIEW

Literature identifies bank failure risk through different aspects. First, the asset risk under the « Weak Fundamentals Hypothesis » (Torna, 2010; Lanine and Vander Vennet, 2006; Wheelock and Wilson, 2000; Peek and Rosengreen, 1997) and second, the transformation risk under the « Liquidity Shortage

Hypothesis» (Diamond and Dybvig, 1983; Gorton and Winton, 2014). In the first hypothesis, capital requirement makes bank more resilient to shocks and protects depositors, creditors and investors. It leads banks to set up more strict criteria for their assets selection (Bolt and Tieman, 2004) and for their risk management (Holmstrom and Tirole (1997).Capital reduces debt ratio, lowering bank risk in presence of deposit insurance (Furlong and Keeley, 1989). Berger and Bouwman (2013) support that capital helps small banks to increase their chance of survival during normal times and during banking or market crises. For medium and large banks, capital improves their survival only during banking crises. Earlier, Jensen and Meckling (1976), in their work on the agency relationship, argued that higher capital reduces agency costs between shareholders and borrowers, improving the financial health of the firm. However, some studies criticize capital regulation. Kim and Santomero (1988) argue that capital requirement reduces the expected return and encourages banks to increase their risk leading to bank failure. Barth, Caprio and Levine (2004) did not find a strong relationship between capital regulation and bank stability in presence of other regulation and supervision. Girod and Bruno (2011) found there is an incentive bias to take excessive risk because of higher cost of capital, which is in contradiction with Basle II goal. Under « Liquidity Shortage Hypothesis », banking is based on confidence that can fail in any time leading to bank run. Indeed, banks, even solvent, can become illiquid due to arbitrary changes in the behavior of depositors or to liquidity shocks. Regulation, such as capital requirements, deposit insurance, lender of last resort or funds injection are proposed to cope with this risk (Diamond and Dybvig, 1983; Diamond and Rajan, 2005; Gorton and Winton, 2014). However, liquidity risk has not received as much attention from regulators until the 2008 crisis, even if Cifuentes *et al.* (2005) state that "liquidity buffers may play a role similar to capital buffers. In some circumstances, liquidity requirements may be more effective than capital buffers in forestalling systemic effects".

In the aftermath of this crisis, Basle III, in its original version of 2010, proposed a set of reforms to raise the quality of capital and to establish liquidity coverage. It proposed two complementary standards of liquidity to strengthen bank resiliency: The Liquidity Coverage Ratio (LCR) that concerns the short-term liquidity risk profile through high quality liquid asset and the Net Stable Funding Ratio (NSFR) that reduces long-term funding risk by requiring enough stable funding to finance their activities. Empirical research examining the link between liquidity risk and bank failure focused on both the sources of funding and the asset composition. Bologna (2013) analyzed the role of funding in explaining U.S. bank failure during the recent crisis through deposit structure. He found that riskier deposits such as brokered and large time deposits are positively linked to bank failure. Cole and White (2012) found that loan portfolio composition (personal versus commercial loans) affects US commercial bank failure and plays a key role in failure prediction while capital ratio loses its predictive power over the long-term. Vazquez and Federico (2012) analyzed the role of both capital and NSFR in explaining U.S. and European banks failure before the 2008 crisis. They showed that liquidity and capital have a complementary role in preserving bank soundness. To our knowledge, few empirical studies explored the impact of Basel III requirements on bank failure. Most of studies used the simple short-term liquidity ratios (Wheelock and Wilson, 2000 and Cole and white, 2012), and did not consider off-balance sheet (Vazquez and Federico, 2012 and Fungacova et al., 2013). Most of these studies did not analyze these issues during the crisis period.

DATA AND METHODOLOGY

The dataset comes from the FDIC annual Call Reports. It consists of an original sample of all US commercial banks, during the 2008-2011 period. We exclude from the sample banks that have zero commercial and industrial loans, and those having deposits held in foreign offices. Information about bank failure comes from the FDIC web site. Based on this information, the dataset consist of 13,105 bank-year observations using financial data for the years 2008, 2009 and 2010, i.e. one year before the failure. There are 478 failed banks during the 2008-2012 period.

Explanatory Variables

Our explanatory variables are at first Basle 3 Tier One, Leverage and long-term liquidity ratio -NSFR-. We use also financial variables such as Asset Quality, Management, Earnings and Liquidity and Control variables. Our expectations of these variables contributions are as follows: Tier One ratio (TO) is Basle 3 capital requirement considered as buffer allowing to absorb any losses incurred on banks assets. We expected that Tier One reduce the likelihood of failure. Leverage (LEV) is a non-risk-based leverage ratio as supplementary to Tier One (BCBS, 2013) which helps to limit the procyclicality of regulatory capital. We expect LEV to reduce bank failure risk. Net Stable Funding Ratio (NSFR) is the long-term liquidity ratio required by Basel III. It needs a classification by category and by maturity for each asset class. This information is not available for loans. We propose two measures of NSFR, one by category (NSFRc), used in our baseline models and one by maturity (NSFRm) used for robustness check. We expect banks with higher NSFR to be less likely to fail.

Asset quality (NPL): nonperforming loans ratio is a measure of loan quality. We expect that higher NPL increase bank failure risk.

Management quality (MNG): we consider cost control as an indirect measure of management quality. The more important are these costs, the more banks are likely to fail. The expected relationship is positive.

Earnings (ROA): the return on average assets measures bank profitability, since it considers all revenues and expenses related to banking. We expect that more profitability reduces bank failure risk.

Liquidity (LIQ): is the ability to satisfy the needs of depositors to withdraw their deposits. It is a measure of short-term liquidity. Thus, less liquid banks are more exposed to failure risk.

Size (LNTA): is the natural logarithm of total assets. It reflects the differences between large and small banks for constraints on credit and diversification. According to "too big to fail" statement, we expect that bank size can be a protection against failure risk.

Empirical Model

We use the panel logistic model to predict bank failure. It is a regression model taking in consideration a dichotomist dependent variable, Fail, which takes the value one for failed banks and zero for non-failed banks. It establishes a relationship between the likelihood that the failure event happens or not and a vector of lagged explanatory variables. The probability that a bank fails at a moment is given by the value of the cumulated logistic distribution evaluated for the data and the estimated parameters. The posteriori probability of failure can be derived from the following equation:

$$Fail_{i,t} = log\left(\left[\frac{P_{i,t}}{\left(1 - P_{i,t}\right)}\right]\right) = \beta_0 + \beta_1 X_{1,t-1} + \beta_2 X_{2,t-1} + \cdots + \beta_n X_{n,t-1}$$

And $P_{i,t} = \frac{1}{1+e^{Fail_{i,t}}}$

Where log is the natural logarithm, P_i is the probability that bank i will fail next period, X_j is the jth independent variable, and β_j is the coefficient of the j th predictor variable. The coefficients measure the effect on the odds of bank failure of a unit change in the corresponding independent variables. Positive coefficients will increase probability of failure, while negative ones will be associated with a decrease in the probability of failure. To test our hypothesis, we estimate the multivariate panel logit models described by equations (1) to (7). We introduce separately Tier One (TO), Leverage (LEV), and NSFR, in equations (1) to (3) respectively to study the relevance of each variable. Equation (4) includes Tier One and Leverage. Equation (5) includes Tier One and NSFR and equation (6) include Leverage and NSFR. In equation (7), we try to mimic the requirements of Basel III by introducing in the same model Tier One, Leverage and NSFR.

$$\text{Fail}_{i,t} = \beta_0 + \beta_1\text{TO}_{i,t-1} + \beta_2\text{NPL}_{i,t-1} + \beta_3\text{ROA}_{i,t-1} + \beta_4 LIQ_{i,t-1} + \beta_5 MNG_{i,t-1} + \beta_6 LNTA_{i,t-1} + \varepsilon_{i,t} \quad (1)$$

$$\text{Fail}_{i,t} = \beta_0 + \beta_1\text{LEV}_{i,t-1} + \beta_2\text{NPL}_{i,t-1} + \beta_3\text{ROA}_{i,t-1} + \beta_4 LIQ_{i,t-1} + \beta_5 MNG_{i,t-1} + \beta_6 LNTA_{i,t-1} + \varepsilon_{i,t} \quad (2)$$

$$\text{Fail}_{i,t} = \beta_0 + \beta_1\text{NSFR}_{i,t-1} + \beta_2\text{NPL}_{i,t-1} + \beta_3\text{ROA}_{i,t-1} + \beta_4 LIQ_{i,t-1} + \beta_5 MNG_{i,t-1} + \beta_6 LNTA_{i,t-1} + \varepsilon_{i,t} \quad (3)$$

$$\text{Fail}_{i,t} = \beta_0 + \beta_1\text{TO}_{i,t-1} + \beta_2\text{NPL}_{i,t-1} + \beta_3\text{ROA}_{i,t-1} + \beta_4 LIQ_{i,t-1} + \beta_5 MNG_{i,t-1} + \beta_6 LNTA_{i,t-1} + \beta_8\text{LEV}_{i,t-1} + \varepsilon_{i,t} \quad (4)$$

$$\text{Fail}_{i,t} = \beta_0 + \beta_1\text{TO}_{i,t-1} + \beta_2\text{NPL}_{i,t-1} + \beta_3\text{ROA}_{i,t-1} + \beta_4 LIQ_{i,t-1} + \beta_5 MNG_{i,t-1} + \beta_6 LNTA_{i,t-1} + \beta_8\text{NSFR}_{i,t-1} + \varepsilon_{i,t} \quad (5)$$

$$\text{Fail}_{i,t} = \beta_0 + \beta_1\text{LEV}_{i,t-1} + \beta_2\text{NPL}_{i,t-1} + \beta_3\text{ROA}_{i,t-1} + \beta_4 LIQ_{i,t-1} + \beta_5 MNG_{i,t-1} + \beta_6 LNTA_{i,t-1} + \beta_8\text{NSFR}_{i,t-1} + \varepsilon_{i,t} \quad (6)$$

$$\text{Fail}_{i,t} = \beta_0 + \beta_1\text{TO}_{i,t-1} + \beta_2\text{NPL}_{i,t-1} + \beta_3\text{ROA}_{i,t-1} + \beta_4 LIQ_{i,t-1} + \beta_5 MNG_{i,t-1} + \beta_6 LNTA_{i,t-1} + \beta_8\text{LEV}_{i,t-1} + \beta_9\text{NSFR}_{i,t-1} + \varepsilon_{i,t} \quad (7)$$

RESULTS AND DISCUSSIONS

We present the means and standard errors for the subsamples of non-failed and failed banks, for the 2008-2010 period in Table 1. We use the t-test to compare the difference in mean between failed and non-failed banks. Not surprisingly, financial variables are statistically significant at the 1% level and present the expected sign. As expected, failing banks have significantly lower NSFR by category. However, the difference in the means for the NSFR by maturity is significant but does not have the correct sign. The correlation matrix for the independent variables shows no excessive correlation between them, except between Tier One (TO) and Leverage (LEV).

Table 1 : Descriptive Statistics for Non-Failed and Failed Banks During the 2008-2010 Period

Variable	**Non-Failed Banks**		**Failed Banks**		**Difference**	**T-Statistic**	
	Mean	**Std. Dev.**	**Mean**	**Std. Dev.**			
TO	13.56	6.38	4.47	3.55	9.09	26.84	***
LEV	1.18	1.48	0.54	0.63	0.63	8.11	***
NSFRc	1.03	0.13	0.98	0.11	0.04	6.17	***
NSFRm	1.74	1.88	2.07	1.8	-0.33	-3.24	***
NPL	2.99	.03	11.13	.48	-8.14	-39.47	***
MNG	89.04	42.14	126.62	128.25	-37.58	-15.03	***
ROA	-0.16	1.95	-5.1	3.26	4.93	46.19	***
LIQ	29.28	17.59	23.71	12.53	5.56	5.94	***
LNTA	12.38	1.14	12.6	1.2	-0.21	-3.49	***

Difference in the means of non-failed and failed banks and t-test for significant differences are presented in the two last columns. Data in Millions of US Dollars. Source Federal Deposit Insurance Corporation.

Empirical Results

We present the regression results of the equations (1- 7) in Table 2. Each target variable introduced separately (1-3) presents the expected contribution with significant coefficients at the 1% level. Higher Tier One ratio, Leverage ratio and NSFR by category reduce bank failure risk. This implies that well capitalized banks and those with sound liability structure are more resilient to failure. These results follow our assumptions, Basel III expectations and former and new literature (Bolt and Tieman, 2004; Berger and Bouwman, 2012, Vazquez and Federico, 2012).

NSFR by category loses its significance with Tier One and Leverage ratios (equations 5, 6 and 7) suggesting the two capital ratios are sufficient in explaining banks failure in contrast to the new Basel III long-term liquidity requirements. Surprisingly, Leverage ratio increases significantly bank failure risk, when introduced with Tier One (equations 4 and 7) in opposition to Basel III hypothesis and to our assumptions. Rugemintwari *et al.* (2012) justify the coexistence of the tier one ratio and the simple leverage ratio by their complementarity when banks underestimate their risk, while it would be superfluous in the absence of cheating and essential when cheating is high. Positive correlation between leverage and bank failure in presence of the tier one ratio suggest that banks underestimate their risk by choosing riskier assets and speculative derivatives. This finding may be the result of the composition of the leverage ratio introduced by Basel III that considers the off-balance sheet items (derivatives, CDS and MBS). These products, designed for coverage objectives, can deviate from their original purpose to become speculative products, increasing bank risks. In fact, the respect of a minimum leverage can be an incentive bias to take excessive risk to compensate the increasing cost of capital. Girod and Bruno (2011) found similar results for capital ratio under Basel II.

Table 2: Baseline Panel Logistic Regression Results for the Entire Sample

Variable	Eq1		Eq2		Eq3		Eq4		Eq5		Eq6		Eq7	
TO	-2.26	***					-2.43	***	-2.22	***			-2.46	***
	(0.18)						(0.19)		(0.18)				(0.19)	
LEV			-1.76	***			1.08	***			-1.75	***	1.15	***
			(0.24)				(0.32)				(0.25)		(0.32)	
NSFRc					-2.56	***			-2.04		-0.2		-3.18	
					(0.9)				(2.25)		(0.95)		(2.33)	
NPL	0.05		0.16	***	0.16	***	0.03		0.05		0.16	***	0.04	
	(0.03)		(0.02)		(0.02)		(0.04)		(0.03)		(0.02)		(0.04)	
ROA	-0.47	***	-0.57	***	-0.6	***	-0.47	***	-0.46	***	-0.57	***	-0.48	***
	(0.08)		(0.06)		(0.05)		(0.09)		(0.08)		(0.06)		(0.09)	
LIQ	0.01		-0.03	***	-0.03	***	0.01		0.01		-0.03	***	0.01	
	(0.01)		(0.01)		(0.01)		(0.01)		(0.01)		(0.01)		(0.01)	
MNG	0.01	***	0		0		0.01	***	0.01	***	0		0.01	***
	(0)		(0)		(0)		(0)		(0)		(0)		(0)	
LNTA	0.36		0.06		0.3	***	0.55	**	0.31		0.06		0.5	**
	(0.22)		(0.08)		(0.08)		(0.23)		(0.22)		(0.08)		(0.24)	
CONS	2.46		-5.25	***	-7.19	***	0.38		4.84		-5.06	***	4.11	
	(2.96)		(1.22)		(1.49)		(3.1)		(3.97)		(1.51)		(4.15)	
ll	-638.2		-968.35		-1023.2		-633.31		-637.86		-968.32		-632.33	
aic	1,292.4		1,952.6		2,062.4		1,284.6		1,293.7		1,954.6		1,284.6	
bic	1,352.2		2,012.4		2,122.2		1,351.8		1,360.9		2,021.9		1,359.4	

In all models, the dependent variable equals one for failing banks, and zero otherwise. The three target variables are Tier One (TO), Leverage (LEV) and NSFR by category.t-statistic between parenthesis.,** and *** indicate statistical significance at the 0.1, 0.05 and 0.01 levels, respectively.*

The coefficients associated to earning are negative and significant at 1% level. This result is consistent with previous studies suggesting that failure risk increases for less profitable banks (Arena, 2008; Lanine and Vennet, 2006; Wheelock and Wilson, 2000). More liquid banks and those with high asset quality are less likely to fail when associated with leverage ratio or with NSFR (equations 2, 3 and 6) in line with Cole and White (2012) findings. However, liquidity and asset quality lose their relevance with Tier one ratio (equations 1, 4, 5 and 7). This result could be justified by the construction of Tier one ratio that considers credit risks, market risks and operational risks. Finally, better management quality reduces bank failure risk. Estimated coefficients for bank size are positive but not statistically significant for all the models.

Robustness Check

To evaluate the importance of alternative criteria for measuring liquidity requirement from Basle 3, we performed an additional estimation to examine the robustness of our results by using an alternative measure of NSFR, i.e. NSFR by maturity (NSFRm). The results in Table 3 show that, surprisingly, NSFRm has a positive and significant impact on bank failure risk. We can explain this result, inconsistent with our assumptions and Basle 3 objectives, by excessive reliance on short-term activity since the denominator of NSFR by maturity, RSF, decreases when the short-term loans are important. Another limit of the NSFRm concerns the problem of accurate credit liquidity evaluation. Long-term credits can be securitized and so become liquid. In absence of precise information about the real credit liquidity, NSFRm is incoherent (Berger and Bouwman, 2009). These results confirm the findings of the baseline model for all the other variables bringing some robustness to our previous results.

Table 3 : Robustness Check by Alternative Measure of Long-Term Liquidity Ratio, (Eq7)

Variable	NSFRc		NSFRm	
TO	-2.46	***	-2.43	***
	(0.19)		(0.2)	
LEV	1.15	***	0.93	***
	(0.32)		(0.34)	
NSFR	-3.18		0.44	***
	(2.33)		(0.14)	
NPL	0.04		0.03	
	(0.04)		(0.04)	
ROA	-0.48	***	-0.45	***
	(0.09)		(0.09)	
LIQ	0.93		1.15	
	(0.73)		(0.99)	
MNG	0.68	***	0.68	***
	(0.25)		(0.24)	
LNTA	0.5	**	0.49	**
	(0.24)		(0.24)	
CONS	4.11		0.36	
	(4.15)		(3.13)	
ll	-632.33		-629.14	
aic	1,284.66		1,278.29	
bic	1,359.41		1,353.03	

In all models, the dependent variable equals one for failing banks, and zero otherwise. The alternative measure of liquidity is NSFR by maturity (NSFRm).t-statistic between parenthesis.,** and *** indicate statistical significance at the 0.1, 0.05 and 0.01 levels, respectively.*

Classification by Size

Size is an important factor in explaining banks failure. The large number of small banks may hide the specificities of medium and large banks. We classify banks by asset size in two groups: small banks with assets less than 1 billion dollars and medium and large banks with assets greater than 1 billion dollars. There are 11 717 small banks and 1 362 medium and large banks. Table 4 presents estimation results by asset size class. For small banks, as expected, the results are similar to the baseline model, in terms of sign and significance. The major result concerns the NSFR that becomes significant at the 10% level reducing banks failure risk as expected by the BCBS. So, small banks, with higher stable funding are less likely to fail. For medium and large banks, only the two capital ratios remain significant. These results show that small banks are more sensitive to their fundamentals (Cole and White, 2012). Bank size variable is significant for medium and large banks, increasing bank failure risk. These banks benefit from advantage of state support (in the form of implicit insurance). This confirms the “too big to fail” hypothesis that incites large banks to take excessive risks.

Table 4 : Logit Regressions by Sub-Samples of Size Category

Variable	Baseline Model, All Size		Small Banks		Medium and Large Banks	
TO	-2.46	***	-2.03	***	-3.9	***
	(0.19)		(0.19)		(0.77)	
LEV	1.15	***	0.88	***	3.17	**
	(0.32)		(0.29)		(1.52)	
NSFRc	-3.18		-4.21	*	3.67	
	(2.33)		(2.28)		(6.66)	
NPL	0.04		0.04		-0.08	
	(0.04)		(0.03)		(0.13)	
ROA	-0.48	***	-0.41	***	-0.46	
	(0.09)		(0.08)		(0.32)	
LIQ	0.93		0.7		0.1	
	(0.73)		(0.69)		(5.98)	
MNG	0.68	***	0.56	**	0.36	
	(0.25)		(0.23)		(1.99)	
LNTA	0.5	**	0.07		2.37	**
	(0.24)		(0.3)		(1.09)	
CONS	4.11		8.78	*	-19.08	
	(4.15)		(4.74)		(16.97)	
ll	-632.33		-552.38		-75.22	
aic	1,284.66		1,124.7		172.44	
bic	1,359.41		1,198.4		229.72	

*Large banks have assets greater than 3 billion dollars, medium banks have assets between 1 and 3 billion dollars and small banks have assets less than 1 billion dollars. In all models, the dependent variable equals one for failing banks, and zero otherwise. t-statistic between parentheses. *,** and *** indicate statistical significance at the 0.1, 0.05 and 0.01 levels, respectively.*

CONCLUSION

Basel III tried to consider broader aspects of banking activity. Its objective is to increase banks resilience through higher capital requirements (Tier One and Leverage), and to reduce banks fragility by introducing two liquidity ratios, LCR and NSFR. We used a panel logit analysis on dataset drawn from the 2008-2010 subprime crisis period for US commercial banks to explore the contribution of these measures in reducing banks failure risk. We found that in presence of the tier one ratio, leverage increases bank failure risk, suggesting that banks underestimate their risk by choosing riskier assets and speculative derivatives. This result points out the importance of considering and supervising the off-balance sheet activity in applying Basel III requirements. In regards to the NSFR, we surprisingly found out that its efficiency depends on the bank size. Banks of smaller size are more sensitive to their fundamentals and to regulatory requirements. Especially, they are more affected (by risk) when there is a lack of stable funding. Banks of larger size, however, benefiting from the implicit insurance, are not affected by these elements. Our study provides some support to Basel III, suggesting to the regulators: first, to choose the Leverage ratio as a complementary constraint for reducing bank regulatory arbitrage. Second, to continue implementing liquidity ratios that provide more buffers against the failure risk for small banks. Large banks, taking advantage from the "too big to fail" and "too big to discipline", are not affected by liquidity measures. So, regulators should strengthen supervision and transparency requirements for these banks.

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ESTIMATING AND ANALYZING THE TECHNICAL EFFICIENCY OF BANKS IN GHANA

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ABSTRACT

This study examined the Technical Efficiency (TE) of commercial banks in Ghana and the determinants of TE in the banking sector. Data Envelopment Analysis (DEA), Random-Effects Tobit regression and Ordinary Least Square (OLS) were the statistical tools used on a sample of 21 banks operating between 2009 and 2013. The results showed that there were more technically inefficient banks in the country than there were technically efficient ones. Another revelation was that, on the average, TE varied directly in proportion to bank size within the two upper quartiles but large banks did not benefit from economy of scales as an edge over small banks. Finally, Gross Domestic Product (GDP) per capita, inflation, credit risk, size and operating cost negatively influenced efficiency while market concentration had a positive influence on efficiency. In our recommendations, we admonished Bank managers to minimize their operating cost and credit risk. Similarly, we recommended that government limit inflation. Finally, we upheld the continuation of the policy of higher capitalization that the Bank of Ghana had been pursuing.

JEL: G21

KEYWORDS: Bank, Technical Efficiency, DEA, Determinants

INTRODUCTION

The role of financial institutions in the development of countries cannot be overemphasized. Efficient Financial institutions are likely to promote economic growth through financial intermediation, investments, employment and taxes to mention but few. Contrarily, inefficiency in the financial sector may lead to credit crunch, recession, financial disintermediation, unemployment etc. It is thus of essence to assess the efficiency of financial institutions in general but particularly that of banks since many studies, including Saka, et al. (2012), identify them as major players in the financial sector.

In Ghana, banks have been rendering financial services prior to independence. After independence, many state-owned banks sprang up with the intention of promoting the drive towards industrialization. Unfortunately, the country traversed severe economic difficulties between 1970 and 1983. This resulted in the accumulation of huge non-performing assets by many banks. Consequently in 1987 the Government of Ghana (GOG) with the assistance of the World Bank, commenced a financial sector reform program to, among other things, establish a sound prudential and regulatory framework for banking operations and liberalize the money and capital markets (Antwi-Asare and Addison, 2000). The liberalization resulted in an increase in the number of commercial banks in the country from 16 in 2000 to 27 in December 2013. However, this also called for extra measures. For example, the Bank of Ghana noted that many of the banks were small in terms of capital base and were therefore highly vulnerable to any adverse but feeble swing in macroeconomic fundamentals. In its opinion, small banks also lacked the capacity to mobilize sizeable equity funds from the international market for investment purposes (Bank of Ghana, 2008). As a result, it resolved to implement a policy that seeks to increase the capital base of banks operating in the country. To obtain a class 1 banking license for instance, the minimum capital requirement was, in 2008, raised to GH¢60 million for all banks and was further augmented to GH¢ 120 million for new entrants in 2013.

This policy corresponds with the requirements of Basel III, which, according to Gual (2011), is an international initiative that provides solutions to some of the inadequacies of the regulatory framework before the banking crisis of 2007 - 2011. In relation to capital requirements, it specifically demands that banks should maintain higher levels of equity in order to be able to deal with potential losses. The initiative also seeks to ensure that financial institutions operate at lower risk levels. There are numerous studies on the efficiency of banks. Many of them are however related to the advanced economies like Canada, United States and member states of the European Union (Allen and Liu (2005), Pasiouras, et al. (2007), Casu and Molyneux (2003), Rozzani and Abdul Rahman (2013) and Wheelock and Wilson (2000)). In Ghana, current researches on the efficiency of banks are very few. Examples thereof include Antwi-Asare and Addison (2000) and Saka, et al. (2012). Given the dynamism of the banking industry (i.e. rapid technological transformation, increased competition, changes in regulatory framework etc.), many of the existing conditions that lead to their findings and recommendations might have changed considerably with time. Therefore, it is likely that the result of these studies may not be relevant contemporarily. That is why it is expedient to conduct new researches on the topic.

This work will be beneficial to bank managers, governments, researchers and bank clients. To bank managers, this study enables them to gauge their efficiency against the best practice and provide them with clues on the determinants of TE in order to guide them on how to keep performance at peak. For researchers, this work contributes to the pool of literature and facilitates comparative surveys or the study of trends in different research works. Burger and Humphrey (1997) is a typical example. To the government, the result of this study provides a feedback on the appropriateness of its policy of increasing the minimum capital requirement. In the event that large banks are better than small ones concerning TE and/or economy of scale, then this result will be a vindication of government's policy. This work seeks to estimate and to analyze the TE of banks in Ghana. In specific terms, it will examine i.) The TE of commercial banks in Ghana ii.) Whether big banks were technically more efficient than small banks, iii.) Whether large banks in Ghana benefited from economy of scale as an advantage over small ones, and iv.) The determinants of TE of commercial banks in Ghana There are five sections in this work. The next section involves a review of theoretical and empirical literature on the efficiency of banks. Section 3 considers issues of methodology whilst section 4 entails an analysis of data and presentation of results therefrom. The final section encompasses conclusions, recommendations and limitations of the study.

LITERATURE REVIEW

The Financial Sector Reform Programs in Ghana

The financial sector of Ghana comprises of insurance companies, banks, savings and loan firms, Ghana Stock Exchange, discount houses, leasing firms, credit unions, foreign exchange bureaus and mortgage financing companies. There had been numerous financial reforms in the country since independence in 1957. However, the most significant among them occurred in the late nineties. According to BOG, 1998/1999 report, cited by Antwi-Asare and Addison (2000) these reforms took place between 1988 and 1990. They sought to among other things: establish a sound prudential and regulatory framework for banking operations, ensure uniform accounting and auditing standards for all banks, and put in place a more effective Banking Supervision Department (BSD) endowed with the requisite personnel and skills to enforce the rules and regulations as well as a code of conduct for the banking sector. Pursuant to this, there was an authorization for the operation of foreign-exchange bureaus in 1988 and there was equally the establishment of the Ghana Stock Exchange in 1989. In that same year, a new banking law was enacted which shunned the imposition of entry or exit restrictions in order to foster competition. The country abandoned Credit ceiling and credit allocation policies in favor of indirect instruments of monetary control. This resulted in the liberalization of interest rates and the relaxation of direct control over bank charges. Moreover, the government recapitalized banks in financial distress by offsetting their non-performing loans with interest-bearing bonds issued by the Non-Performing Asset Recovery Trust (NPART). The latter was responsible for the retrieval of these loans. The change also led to the replacement of top management of many state-owned banks and the reconstitution of many Board of Directors.

The reforms led to a rapid increase in the number of banks in the sector (from 16 in 2000 to 27 in 2007). A development that compelled BOG to institute further measures because of the conviction that the numerous banks springing up had low capital base and were unable to provide high levels of lending, notably on the international platform. Again, the Central Bank was worried that small banks were highly susceptible to the slightest adverse swing in macroeconomic fundamentals. More importantly, BOG considered the domestic commercial banks as lacking the requisite capacity to enable them raise mammoth equity finance through the capital market. Consequently, over the past few years, BOG has been pushing for larger minimum capital for financial institutions within its jurisdiction. For example, in order to obtain a Class 1 license, commercial banks had their minimum capital requirement raised to GH¢60 million while banks in operational existence were obliged to meet that same level latest by the end of December 2009. Local banks were provided with an extended period to enable them increase their capital to GH¢ 25 million latest by the end of 2010 and to further augment it to GH¢60 million by 2012 (Bank of Ghana, 2008). This minimum capital requirement was further raised to GH¢ 120 million for new entrants seeking the class 1 banking license while existing class 1 banks were advised to take steps to enhance their capital in line with their business strategy and risk profile to avoid requesting for single obligor exposure waivers (Bank of Ghana, 2013).

Efficiency in Literature

The notion of efficiency crisscrosses numerous disciplines such as Economics, Physics and Medicine. Heyne (2008) explained that efficiency is the rapport between ends and means. A condition is not efficient if it could not attained its targeted end with fewer resources or if the resources used yielded less than the expected results. Farrell (1957) and De Borge, et al. (1994) added that a firm achieves efficiency when it succeeds in turning out the largest feasible output from the inputted resources. They then distinguished between price efficiency and TE. The former refers to an entity's accomplishment in selecting the best combination of inputs while the latter gauges its achievement in turning out the highest level of output from a given collection of inputs Economists distinguish among allocative efficiency, productive efficiency, TE, x-efficiency, dynamic efficiency, and static efficiency. This study focuses on TE but details on other types are obtainable from Kolasky and Dick (2003) and Economics Online (2014). Maidamisa, et al. (2012) mentioned that it is possible to divide TE into pure TE and scale efficiency. Pure TE measures how a producer utilizes its resources under exogenous environments and scale efficiency is the percentage of TE to pure TE. How do we measure TE?

In 1951, Koopmans and Debreu initiated the efforts to measure efficiency and Farrell made his input in 1957. Thereafter, researchers developed numerous techniques on the measurement of efficiency. Mandl, et al. (2008) mentioned that TE improvements are drifts towards the Production Possibility Frontier (PPF), which other literature also refer to as efficiency frontier. However, not all forms of TE have economic significance because TE ignores costs and benefits resulting from the optimum arrangement of inputs. This implies that the achievement of a high level of TE is not economically sensible so long as alternative permutation of inputs would lead to higher outputs. The measurement of allocative efficiency addresses the weaknesses of TE. Allocative efficiency examines the connection between optimum arrangements of inputs taking on board not only costs and benefits, but also the output achieved. In the next paragraph, we will examine the techniques used in estimating the efficiency frontier.

Parametric and non-parametric approaches are the two major techniques used in estimating the efficiency frontier. . Mandl, et al. (2008) maintained that parametric frontier functions require an assumption of a particular form of association between input and output. There is equally an assumption that data or population under study has a specific structure (i.e. normal distribution, F-distribution or Chi-Square distribution). Contrarily, a non-parametric procedure has little or no a priori assumption. In order words, presumption about data or population having a specific characteristic structure is almost inexistent. Consequently, non-parametric procedures tend to be flexible and susceptible to several alternative formulations. Berger and Humphrey (1997) concluded that efficiency estimates from non-parametric studies are akin to those from parametric frontier model. However, the nonparametric methods normally turn out lower average efficiency scores and appear to have larger variations than the results of parametric models. Murillo-Zamorano and Vega-Cervara (2000) also distinguished between deterministic and stochastic models. Deterministic models envelop all the observations, with reference to the gap separating the observed output and the highest possible output. On the other hand, stochastic approaches facilitate the distinction between TE and statistical noise.

For example, the Free Disposable Hull (FDH) is a deterministic and non-parametric tool for estimating productive efficiency while the DEA is a stochastic and non-parametric approach. Examples of parametric frontier approaches are the Stochastic Frontier Approach (SFA), Distribution Free Approach (DFA), and the Thick Frontier Approach (TFA) (Berger and Humphrey, 1997). Within the banking sector, literature on the composition of input and output, used in measuring efficiency focuses mainly on two models, which are the intermediation approach and the production approach. The production approach, as explained by Berger and Humphrey (1997), considers financial institutions as producing services for account holders. These services include, but not limited to processing of loan applications, credit reports, honoring of checks or other payment instruments. Consequently, with the production approach, the inputs considered are only physical inputs such as labor and capital as well as the costs thereof. On the other hand, the measurement of output encompasses factors such as the numerical size and category of documents or transactions processed over a time interval. However, due to the challenges in obtaining vivid details of this data, alternatives such as the number of deposit or loan accounts serviced are used. Now it is important to consider the determinants of efficiency in the banking sector.

Previous works identified profitability ratio, average capital ratio, bank size, market share, credit risk, cost of operation, foreign ownership, market concentration and the security market as key determinants of bank efficiency. Moreover, other literature added macroeconomic variables such as inflation, GDP per capita and interest rate spread (Adjei-Frimpong, et al. (2014), Casu and Molyneux (2003), Pasiouras, et al. (2007), Ayadi (2013), Eriki and Osifo (2015) and Alrafadi, et al. (2014)). Whereas profitability ratio, average capital ratio, bank size, credit risk, cost of operation are accounting ratios and are easily computed from financial statements, other variables like inflation, GDP per capita and interest rate spread, market concentration, market share are more inclined towards Economics. This study uses a combination of both variables.

Empirical Literature

A study of 130 works on the efficiency of financial institutions across twenty one countries, conducted by Berger and Humphrey (1997), found that deregulation of financial institutions as well as mergers and acquisitions can either ameliorate or aggravate efficiency depending upon the prevailing conditions in the industry before to deregulation. The case of numerous states showed that deregulation resulted in fast expansion of branches, massive asset increases, bank failures, and reduced efficiency. Similarly, mergers and acquisitions aggravated the performance of the merged firms compared with the individual institutions existing on their own. Vegesna and Dash (2014) employed the DEA approach to examine the efficiency of banks in India between 2005 and 2011 comparing public sector banks to their counterpart in the private sector. They found that public sector banks in India were more efficient than private banks. In a study of Canadian Banks from 1983 to 2003, Allen and Liu (2005) found that larger banks appeared to be more cost-efficient than smaller banks. They could enjoy cost savings of 6% to 20% by increasing their scale of production. Using SFA and multiple linear regression to arrive at their conclusion on a research examining the determinants of bank efficiency among conventional banks and Islamic banks in Malaysia, Rozzani and

Abdul Rahman (2013) stated that, there was no much difference between the level of profit efficiency between conventional banks and Islamic banks. Banks in the country had an overall efficiency falling below 50%. There was also a mammoth positive association between bank size and conventional bank efficiency but no such association existed between bank size and Islamic bank efficiency.

Contrarily, we found an essential negative correlation between operational cost and efficiency for both conventional and Islamic banks. Finally, there was no significant negative relationship between credit risk and Islamic bank efficiency. Nonetheless, conventional banks had an important negative relationship between credit risk and efficiency. Casu and Molyneux (2003) employed DEA and Tobit regression to examine the determinants of European bank efficiency. The results proved that average efficiency was low although it improved over time. They also mentioned that the key factors distinguishing among bank efficiency levels across the continent were country-specific factors of banking technologies. There was however, little to show that average capital ratio and return on average equity explained variations in bank efficiency levels. Similarly, Pasiouras, et al. (2007) employed DEA and Tobit regression to analyses the determinants of bank efficiency in Greek. Their result showed the possibility of improving average cost efficiency by 17.7%. They equally found size had a positive impact while GDP per capita had negatively and materially affected all measures of efficiency. Unemployment rate had a substantial negative impact on both technical and cost efficiencies but not on allocative efficiency.

Ayadi (2014) used DEA to analyses the technical efficiency of banks in Tunisia and their determinants from 2000 to 2011. He estimated technical efficiency of banks in Tunisia to be 57.1%; pure technical efficiency and scale efficiency were 64.7% and 86.9% respectively. He concluded that high bank capitalization had a positive impact on technical efficiency of the banks while market share in terms of deposit of the banks adversely affected their technical efficiency. He added that private banks were more efficient that their public counterparts but size had no bearing on technical efficiency considering the sample studied. Alrafadi, et al. (2014) made use of DEA and Tobit regression to ascertain the level of efficiencies of Libyan banks from 2004 to 2010 and their determinants. They held that, on one hand Libyan banks had a mean TE of 59.3% and could therefore minimize input by 40.7% at the same output level. On the other hand, factors like return on assets, size, government ownership and capital adequacy influenced the efficiency of bank positively. Making use of DEA and multiple regression, Kamau and Were, (2013) looked at the forces that drove the high performance in the Kenyan banking sector between 1997 and 2011 by analyzing the relationship structure, performance and bank efficiency. They concluded that structure and/or the collusion power of industry players were the major factor propelling high performance rather than efficiency. In a study of the impact of the 2008 - 2009, global financial crisis on large South African banks, Erasmus and Makina (2014), used DEA to arrive at the conclusion that most of the big banks were efficient and therefore the crisis had no adverse influence on them.

Eriki and Osifo (2015) found out the determinant of performance efficiency of a sample of 19 Nigerian banks using DEA. They reached the conclusion that small and medium banks were better than huge banks in terms of performance efficiency. They equally found that board independence and bank ownership had negative association with Nigerian bank efficiency but bank size and age had a positive influence on efficiency. These findings appear inconsistent because if small and medium banks outperformed huge banks one would expect size to have a negative relation with efficiency. Mawutor and Fred (2015) also employed Return on Assets (ROA) ratio and panel regression model to assess the efficiency and profitability of listed banks in Ghana from 2006 to 2011 and stated that productivity, loan and credit risk had a significant negative effect on profitability but liquidly had a significant positive effect on it. Bank size, according to them had no influence on productivity. Adjei-Frimpong, et al. (2014) employed DEA, fixed effect model and the Generalized Method of Moment (GMM) to study the cost efficiency of the Ghanaian banking industry. The results suggest that banks in Ghana were inefficient. Moreover, the fact that bank size had no relation with cost efficiency led the authors to state that larger banks in Ghana had no cost advantages over smaller ones. Other results are that, the ratio for provision for loan loss had no effect on bank efficiency but GDP growth rate negatively affected bank cost efficiency.

Tetteh (2014) used ROA ratio as well as panel regression model to measure the performance of local and foreign banks in Ghana between 2003 and 2012. He observed that generally, foreign banks outperformed their local competitors. Furthermore, market experience, size and ownership had substantial negative impact on bank performance whilst success in the market, and interest income were positively significant in influencing performance. The implication is that lengthy years of experience and large sizes reduced bank performance. Saka, et al. (2012), used DEA and Tobit regression to study the determinants of Technical Efficiency (TE) of Ghanaian banking industry and the impacts of the entry of foreign banks. They found that the entry of foreign banks had a positive impact on the technical efficiencies of domestic banks although the TE of new foreign entrants was lower than the TE of new local entrants. They attributed the lower TE experienced by foreign entrants to high start-up costs. They also stated that foreign ownership, return on assets and inflation had a significant positive influence on TE while market concentration, loan ratio and capitalization ratio had a negative relation with TE.

Oteng-Abayie, et al. (2011) relied on the maximum likelihood estimation to study the determinant of efficiency of microfinance institutions. They concluded that variations in technical expertise (in both training and portfolio quality) and management practices accounted for inefficiencies in the sector. Buchs and Mathisen, (2005), using Panzar and Rosse Model, studied the competition and the level of efficiency in the Ghanaian banking industry and concluded that scale matters significantly in the Ghanaian banking system. They stated that bank size could constitute a hindrance to market entry. They added that excessive domestic borrowing by government has not only contributed to inefficiency in the banking system but also limited competition between banks because it was profitable and safer for banks to hold government securities. Other impacts of government domestic borrowing are the crowding out of the private sector in accessing bank loans and high interest rates. Finally, they recommended that government should strive to achieve effective fiscal adjustment in order to widen and increase the efficiency in the banking sector. This, they argued would reduce the overreliance of banks upon Treasury bill investment with high return but low-risk. Consequently, this would generate stiffer competition that would in turn propelled the banks to look for ways and means to widen their clientele in order to augment revenue.

DATA AND METHODOLOGY

We obtained the annual financial statements and macro-economic variables from the Bank of Ghana and the website of the Ghana Statistical Service. Efficiency Measurement System version 1.3.0 facilitated DEA analysis whilst Stata/MP 13.0 enabled the computation of the regressions. We use the following input and output variables as shown in Table 1 for the intermediation and production approaches.

Table 1: Variables Used for DEA Computations

Production Approach Variables		Intermediation Approach Variables	
Inputs	Outputs	Inputs	Outputs
Fixed Assets	Investments	Fixed Assets	Interest Income
Non-Interest Expenses	Loans	Non-Interest Expenses	Deposit with other banks
Staff Cost	Deposit with Central Bank	Staff Cost	Non-Interest Income
Interest Expenses	Deposit with other banks		
	Non-Interest Income		

This table shows the composition of input and output variables that enable the computation of DEA for both the intermediation approach and the production approach. In the production approach, the inputs considered are only physical inputs such as labor and capital as well as the costs thereof. On the other hand, the measurement of output encompasses factors such as the numerical size and category of documents or transactions processed over a time interval. However, due to the challenges in obtaining vivid details of this data, alternatives such as the number of deposit or loan accounts serviced are used. Contrarily physical inputs are not necessarily the inputs of the intermediation approach.

The population of banks in the country in 2013 was 27 but we sampled 21 for this work. We equally used a balanced panel on cross-sectional annual data spanning from 2009 to 2013. The adoption of International Financial Reporting Standard (IFRS) by banks in Ghana in 2007 influenced the choice of this time interval. Therefore commencing with 2009 enabled this work to study financial statements prepared according to IFRS after two years of experience with IFRS. Table 2 contains the list of banks in the sample.

Table 2: List of Banks in the Sample

Abbreviation	Official Name
ACCESS	Access Bank
ADB	Agriculture Development Bank
BARODA	Bank of Baroda
BOA	Bank of Africa
BSIC	Sahel Sahara Bank
CAL	Cal Bank
ECO	Eco Bank
FABL	First Atlantic Bank
FBL	Fidelity Bank
GCB	Ghana Commercial Bank
GTB	Guaranty Trust Bank
HFC	HFC Bank
ICB	International Commercial Bank (now FBN Bank)
PBL	Prudential Bank Ltd
SCB	Standard Chartered Bank
SG-SSB	SG-SSB (now Society General Ghana Ltd)
STANBIC	Stanbic Bank
UBA	United Bank for Africa
UNIBANK	Unibank Ghana Ltd
UTB	UT Bank
ZENITH	Zenith Bank Ghana Ltd

This table indicates the official name of banks (and their respective abbreviation) used in the sample. It contains 21 banks derived from a population of 27 banks as of December 2013. All the 21 banks had their Financial Statements prepared according to IFRS. This enhances comparability and could have a positive impact on the results of the analysis.

DEA Technique

The “two-step” procedure was applied in this work using DEA (both production and intermediation approaches) at the first stage and Tobit regression and OLS at the second stage. DEA is a stochastic non-parametric model, which is flexible and susceptible to several alternative formulations. This is because there is virtually no assumption of a specific functional form of relationship between input and output. Moreover, the efficiency frontier was determined using the whole sample. Another strength of DEA is that, it enables a distinction between TE and statistical noise. This feature is relevant to the achievement of objectives of this work. DEA also allows the division of TE into scale efficiency and pure technical efficiency. One can categorize DEA model into Constant Return to Scale (CRS) and Variable Return to Scale (VRS). Charnes, et al. (1978) developed the first DEA model and named it CCR (Charnes, Cooper and Rhodes) model. Among the assumptions, underlying CCR is constant return to scale. Others are strong disposability of inputs and outputs and convexity of the set of feasible input-output combinations. CRS means producers are linearly able to scale the inputs and outputs without increasing or decreasing efficiency (DEA Home Page, 1996). For further clarification, strong disposability signifies the likelihood of varying the quantity of an input or output factor without varying the output produced. CRS relies on the assumption that all DMUs under consideration operate at an optimal scale. In reality however, this assumption rarely materializes due to factors like externalities, financial challenges, imperfect competition, to mention but few. Consequently, the use of CCR model brings about a misleading measure of technical efficiency because technical efficiency scores reported in the midst of these constraints suffer from distortions due to the negligence of scale efficiencies. In an attempt to rectify this shortcoming, Banker, et al. (1984) developed VRS by introducing one additional constraint to those of CRS. This constraint only ensures that the comparison of each DMU is solely against others of similar size. Another name for VRS is BCC (Banker, Charnes and Cooper) model. Mathematically, for any *d* DMUs, producing each *p* products after using *m* input factors, the following Linear Programming (LP) expresses the input-oriented DEA model: in

Minimize θ_i subject to (1)

$Y\lambda_i - y_i \geq 0$ (2)

$\theta_i x_i - X^i\lambda_i \geq 0$ (3)

$\Sigma\lambda_i \geq 0$ (DEA – CCR) (4)

$\Sigma\lambda_i = 1$ (DEA – BCC) (5)

Where: θ_i is a scalar, which stands for the efficiency of the i^{th} firm. λ_i is a percentage of other producers used to generate the virtual producer. *X* is the d×m input matrix represented by the vector x_i for the i^{th} firm. *Y* is the d×p output matrix represented by the vector y_i for the i^{th} firm.

The efficiency score for each firm is determined by solving the LP for each of the d firms. Equations 1 to 4 represent CCR whilst equations 1 to 5 represent BCC. With the first constraint, the virtual DMU is compelled to produce at least as many output as the real DMUs under study. The second constraint also forces the virtual DMU to equal or less input that the real DMUs. In this work, we shall follow the above approach although there are other formulations such as the ratio approach or the dual approach. To compare the efficiencies according to bank size, we grouped banks in the sample into four quartiles based of their average total assets and we named them as "largest banks", "large bank", "medium banks" and "small banks". Given that there were twenty-one (21) banks in the sample, we placed Six (6) banks into the large banks quartile. The other three (3) quartiles have five (5) banks each but on a year-by-year average, the banks that made up each quartile varied. Based on the overall average efficiency of each quartile over the period, we compare the level of efficiencies by size.

Regression Models

Katchova (2013) mentioned that Tobit regression is an example of limited dependent variable models. Gregorian and Monole, (2002) supports the use of Tobit model in the second stage of the two-stage procedure because according to them, efficiency estimate of DEA is censored and limited to variation between the range of 0 and 1. They added that Tobit models generate consistent estimates of regression coefficient. In Tobit regression, there is a restriction on the dependent variable either at lower level (lower limit) or the upper level (upper limit) or both. If the parameter of the dependent variable falls below the lower limit, then it is restricted to the lower limit. Similarly, if its parameter exceeds the upper limit, then we confine it to the upper limit. Below is the mathematical representation of these statements:

$$y = \begin{cases} y^* if\ y^* > L \\ L\ if\ y^* \leq L \end{cases} \tag{6}$$

$$y = \begin{cases} y^* if\ y^* < U \\ U\ if\ y^* \geq L \end{cases} \tag{7}$$

Where: U is the upper limit, L the lower limit, y the actual value of the dependent variable, and y^*latent variable. Equations 4 to 6 represent the empirical regression models.

$$VRS\ TE = \beta_0 + \beta_1 SIZE + \beta_2 OPRC + \beta_3 CRED + \beta_4 OWNR + \beta_5 INFL + \beta_6 MKTC + \beta_7 GDPPC + \varepsilon \tag{8}$$

$$CRS\ TE = \beta_0 + \beta_1 SIZE + \beta_2 OPRC + \beta_3 CRED + \beta_4 OWNR + \beta_5 INFL + \beta_6 MKTC + \beta_7 GDPPC + \varepsilon \quad (9)$$
$$SE = \beta_0 + \beta_1 SIZE + \beta_2 OPRC + \beta_3 CRED + \beta_4 OWNR + \beta_5 INFL + \beta_6 MKTC + \beta_7 GDPPC + \varepsilon \quad (10)$$

Another name for Tobit regression is Censored Normal Regression. In the absence of censoring of observations, then Tobit regression is the same as the Ordinary Least Square (OLS) regression. The equation for the OLS regression is similar to equations 4 to 6.

Where: VRS TE = Variable Return to Scale Technical Efficiency, *CRS TE* = Constant Return to Scale Technical Efficiency, *SE* = Scale Efficiency, *SIZE* = Bank Size, *OPRC* = Operation Cost, *CRED* = Credit Risk, *OWNR* = Ownership, *INFL* = Inflation, *MKTC* = Market Concentration, *GDPPC* = Per Capita Gross Domestic Product, β_0 = Constant coefficient for the regression model, β_1 to β_7 are Coefficients for the dependent variables. The following shows the estimation of the variables:

VRS TE = Output of DEA – BCC

CRS TE = Output of DEA – CCR

$SE = \frac{DEA-CCR}{DEA-BBC}$

$SIZE$ = ln (Total assets)

$OPRC = \frac{Operating\ Cost}{Operating\ Income}$

$CRED = \frac{Total\ loan}{Total\ assets}$

$GDPPC = \frac{GDP}{Midyear\ Population}$

$\mathrm{M}S_i = \frac{Bank\ Total\ Asset}{Industry\ Total\ assets}$

OWNR = Dummy variable (foreign = 1, local = 0)

INFL = Annual average inflation

$MKTC = \sum_{i=1}^{n}(\mathrm{M}S_i)^2$ where $\mathrm{M}S_i$ is Market Share

RESULTS AND DISCUSSION

Results of DEA Production Approach

The production approach averagely showed that three banks were technically efficient within the period based on DEA – CCR model (Table 3) as against eight technically efficient banks based on DEA – BCC model (Table 4). The overall mean efficiency stood at 63.89% for DEA – CCR and 80.48% for DEA – BCC. In conclusion, the production approach estimated that, averagely 61.90% to 85.71% of banks were technically inefficient between 2009 and 2013. Furthermore, the average bank in the country was 63.89% to 80.48% as efficient as the best banks.

Table 3: Summary of Results of DEA Production Approach (CCR –Model)

	2013	2012	2011	2010	2009	OVERALL AVERAGE
Number of DMUs	21	21	21	21	21	21
Number of Efficient DMUs	2	5	5	2	2	3
Number of Inefficient DMUs	19	16	16	19	19	18
Average Efficiency (M)	59.86%	67.92%	68.75%	62.70%	60.24%	63.89%
Average Inefficiency (1-M)	40.14%	32.08%	31.25%	37.30%	39.76%	36.11%
	2013	**2012**	**2011**	**2010**	**2009**	**OVERALL AVERAGE**
Standard Deviation (Θ)	0.2514	0.2480	0.2365	0.2283	0.2286	0.238
Interval I (M-Θ ; M+Θ)	34.72% - 85.01%	43.12% - 92.72%	45.10% - 92.39%	39.87% - 85.53%	37.38% - 83.1%	40.04% - 87.75%
% of DMUs in I	61.90%	57.14%	52.38%	61.90%	61.90%	59.05%

Source: Author's Computation. This table shows the results of the DEA Production Approach (CCR – Model). It also points out the number of technically efficient banks, the average efficiency and standard deviation. The overall mean efficiency stood at 63.89% for DEA – CCR. The production approach estimated that, averagely 85.71% of banks were technically inefficient between 2009 and 2013. Moreover, the average bank in the country was 63.89% as efficient as the best banks.

Table 4: Summary of Results of Production Approach (BCC – Model)

	2013	2012	2011	2010	2009	Overall Average
Number of DMUs	21	21	21	21	21	21
Number of Efficient DMUs	9	11	10	7	5	8
Number of Inefficient DMUs	12	10	11	14	16	13
Average Efficiency (M)	71.56%	86.81%	86.54%	80.47%	77.03%	80.48%
Average Inefficiency (1-M)	28.44%	13.19%	13.46%	19.53%	22.97%	19.52%
Standard Deviation (Θ)	0.2805	0.1798	0.1668	0.2104	0.2176	0.2110
Interval (I) (M-Θ ; M+Θ)	43.51% - 99.60%	68.83% - 104.79%	69.86% - 103.22%	59.43% - 101.51%	55.27% - 98.80%	59.38% - 101.58%
% of DMUs in I	33.33%	76.19%	80.95%	85.71%	80.95%	71%

Source: Author's Computation. This table shows the results of the DEA Production Approach (BCC – Model). It also points out the number of technically efficient banks, the average efficiency and standard deviation. The overall mean efficiency stood at 80.48% for DEA – BCC. The intermediation approach estimated that, averagely 61.90% of banks were technically inefficient between 2009 and 2013. Moreover, the average bank in the country was 80.48% as efficient as the best banks.

The interpretation of the results of intermediation approach is analogous to that of the production approach. Therefore, for the avoidance of repetition, the focus will be on comparing the number of efficient banks and the overall average efficiencies given by the production approach to those of the intermediation approach.

Comparison of Results of Intermediation and Production Approach

From Panel A of Table 5, it is clear that the intermediation approach produced a greater number of efficient banks and a higher overall efficiency score than the results estimated by the production approach. However, a common feature of the results of both CCR and BC is that the number of efficient banks is less than 50% of the number of banks in the sample. The models estimated average efficiency of banks to be between 63.89% and 80.48% whilst the percentage of efficient banks, given the sample size of 21, varied between 14.29% and 42.86%.

Table 5: Comparison of Efficiency Results

	DEA			
Measure	**Production Approach CCR**	**Production Approach BCC**	**Intermediation Approach CCR**	**Intermediation Approach BCC**
Panel A: Efficiency of Banks				
Number of efficient banks	3	8	8	9
Average efficiency	63.8%	80.48%	67.38%	77.99%
Panel B: Technical Efficiency by Bank Size				
Largest banks	71.98%	89.71%	84.4%	90.50%
Large banks	67.18%	83.27%	65.18%	77.24%
Medium banks	66.44%	79.65%	51.88%	70.06%
Small banks	48.36%	65.39%	64.63%	70.91%
Panel C: Number of Technically Efficient Banks by Size				
Largest banks	1	4	4	5
Large banks	1	2	2	2
Medium banks	1	1	1	1
	DEA			
Measure	**Production Approach CCR**	**Production Approach BCC**	**Intermediation Approach CCR**	**Intermediation Approach BCC**
Panel C: Number of Technically Efficient Banks by Size				
Small banks	0	1	1	1
Panel D: Scale Efficiency by bank Size				
Largest banks	78.18%		91.68%	
Large banks	83.23%		63.99%	
Medium banks	80.43%		79.45%	
Small banks	78.05%		81.91%	

This table compares the results of production approach to those of the intermediation approach of the DEA –CCR model and DEA-BCC model. It contains five panels. Panel A estimates the average efficiency of banks and the number of efficiency banks regardless of bank size. Panel B estimates the TE by bank sizes at classifying them into largest bank, large bank, medium banks and small banks. Panel C then examines the number of technically efficient banks per each of the four categories. Finally, Panel D displays the scale efficiency according to bank size. Throughout the panels, the BCC models recorded the highest results compared to the CCR models. Panel D shows that average efficiency varied in proportion to bank sizes only in the upper quartiles.

Technical Efficiency by Bank Size

Panel B of Table 5 shows a decrease in average TE as one moves from largest banks to small banks. The results for both DEA – CCR and DEA – BCC of the production approach follow the same pattern. This implies that the larger the bank the better its technical efficiency. However, observing the intermediation of the same panel, one experiences a slight change in the above-mentioned trend. The results of both DEA – CCR and DEA - BCC of the intermediation approach indicate that largest banks were technically more efficient than all the banks in the other three groupings. Similarly, large banks were technically more efficient than medium banks and small banks. Nevertheless, unlike the production approach, the intermediation approach reveals that small banks were technically more efficient than medium banks. The margin by which small banks outperformed medium banks was very immaterial from the DEA – BCC point of view but significant from the DEA – CCR point of view.

To find out how many banks were technically efficient in each of the four groupings, we carried out the analysis in Panel C of Table 5. The results of the production approach of DEA – CCR model indicate that only one bank was efficient in each of the following categories: largest banks category, large bank category and medium bank category. No bank was efficient in the small bank category. Compared to the DEA – BCC model of the production approach, four banks were efficient in the largest bank category and two in the large banks category. Medium banks and small banks had one efficient banks each. In addition, Panel C of Table 5 shows the results of the intermediation approach of DEA. Both the CCR and the BCC approaches indicate that largest banks had the highest number of efficient banks (four for CCR and five for BCC). Then, large banks followed with two efficient banks each for both CCR and BCC. Medium and small banks followed large banks and had the same number of efficient banks (one efficient bank for both CCR and BCC). The conclusion one could draw is that, generally, largest banks had the highest number of efficient banks and large banks followed. Even where the number of efficient banks was the same among largest banks and the other category of banks (the case of intermediation approach, DEA – CCR model), largest banks had the highest relative efficiency score (Panel B of Table 5).

Scale Efficiency by Bank Size

Panel D of Table 5 shows the result of scale efficiency by bank size. One does not see any clear pattern that may lead to any meaningful conclusion. The results of the production approach indicate that, on the average, large banks had the highest scale efficiency. In order of significance, medium banks, largest banks, and small banks succeeded. On the other hand, the results of the intermediation approach show that, averagely, largest banks were the most efficient. Then small banks and medium banks followed sequentially. One could therefore infer that large banks did not enjoy economy of scales as an edge over small banks.

Random-Effects Tobit Regression

Table 6 summarizes the result of the Tobit regression using CRS, VRS and SE as response variables. Two different models are regressed on each response variables in order to deal with multicollinearity. The symbol ** represents statistical significance at 5% level. The meaning of this symbol applies to all other tables on regression. We discuss the result as follows. From Model 1 to Model 4, OPRC has a significant negative relation with CRS TE, and VRS TE. Inflation and GPDPC have a significant negative influence on CRS TE in Model 1 and VRS TE in Model 3 and Model 4. GDPPC is also significantly associated with VRS TE in Model 3 and Model 4. CRED has a substantially negative association with CRS TE in model 1 while MKTC influences VRS TE negatively in Model 3. OWNR and SIZE have no association with TE and no explanatory variable influences SE. In brief, CRED, OPR, INFL, GDPPC and MKTC have a significant negative influence on VRS TE and CRS TE but OWNR and SIZE have no influence on them. No independent variables affects SE.

OLS Regression

For each model of OLS, we conducted Hausman specification test to determine which model between the fixed-effect model and random effect would be appropriate. Table 7 contains the results of the test. With the exception of Model 3 and Model 4, the Hausman test favored the application of random effect. Therefore, in addition to analyzing the result of random-effect models in Table 7, we also analyze Model 3 and Model 4 using the fixed-effect. The following are the result of the Random-Effect OLS Regression. OPRC has a significant negative relation with CRS TE and VRS TE in models 1 to 4, but has no association with SE (Table 7). SIZE is the only variable that significantly relates to SE (Model 6). This relation is positive but SIZE has no important influence on any other response variable. In Model 3 and Model 4, INFL and GDPPC have negative influence on VRS TE. MKTC also influences VRS TE positively.

In summary, OPRC negatively affects CRS TE while OPRC, INFL and GDPPC negatively influence VRS TE. MKTC also influences VRS TE but in a positive direction. Of all the explanatory variables, only SIZE relates to SE and that relation is positive. As mentioned earlier, we recomputed Model 3 and Model 4 (in Table 8) using the fixed-effect model because the Hausman test favored it. The results indicate that INFL, GDPPC and MKTC influence VRS TE. However, whereas the relations of INFL and GDPPC with VRS TE are negative, that of MKTC with VRS is positive. In summary, Tobit regression, random-effect OLS regression and fixed-effect OLS regression show that INFL, GDPPC and MKTC influence TE. Whereas the influence of INFL and GDPPC are negative, that of MKTC is positive. This implies that increases in INFL and GDPPC lead to a decrease in efficiency whilst an increase in MKTC leads to improvement in efficiency. With regards to company-specific variables, both Tobit regression and random-effect OLS regression points at OPRC as having a negative impact on TE; implying that increases in OPRC would hinder TE. According to random-effect OLS, SIZE negatively influences SE whilst Tobit indicates that CRED has a negative impact on TE. By inference, SE tends to suffer as banks get bigger and increases in CRED cause TE to decline. None of the regression models establishes a substantial association between OWNR and efficiency.

Table 6: Random-Effects Tobit Regression Analysis – Determinants of Bank Efficiency

Dependent Variables	**(1) CRS TE**		**(2) CRS TE**		**(3) VRS TE**		**(4) VRS TE**		**(5) SE**		**(6) SE**	
Independent Variables	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***
CRED	-0.4997	0.042**	-0.3463	0.077	0.0268	0.882	-0.0204	0.911	0.0943	0.587	0.1744	0.391
OWNR	0.1044	0.187	0.0960	0.166	0.0153	0.813	0.0112	0.864	0.0911	0.094	0.1139	0.074
SIZE	0.0201	0.292	0.0209	0.218	0.0022	0.889	0.0051	0.754	0.0256	0.095	0.0252	0.15
OPRC	-0.1570	0.038**	-0.1534	0.022**	-0.1883	0.003**	-0.1818	0.006**	-0.0414	0.486	-0.0531	0.435
INFL	-3.3076	0.022**	-0.1090	0.154	-3.9239	0.001**	-1.4883	0.002**	0.6153	0.218		
GDPPC	-0.1435	0.103			-0.2043	0.004**	-0.2140	0.004**	0.0972	0.188	0.0539	0.434
MKTC	29.762	0.099	-7.4539	0.253	33.578	0.022**						
cons	0.3155	0.797	1.8559	0.036	0.7723	0.437	2.6650	0	-0.6634	0.27	-0.2988	0.561
Prob > chi2	0.0021		0.0095		0.0092		0.0051		0.0488		0.0499	
Observation	105		105		105		105		105		105	

*This table summarizes the result of the Random-Effect Tobit Regression using CRS, VRS and SE as response variables. Two different models are regressed on each response variable in order to deal with multicollinearity. The symbol ** represents statistical significance at 5% level. CRED, OPR, INFL, GDPPC and MKTC have a significant negative influence on VRS TE and CRS TE but OWNR and SIZE have no influence on them. No independent variables affects SE.*

Table 7: Random-Effects OLS Regression Analysis – Determinants of Bank Efficiency

Dependent Variables	**(1) CRS TE**		**(2) CRS TE**		**(3) VRS TE**		**(4) VRS TE**		**(5) SE**		**(6) SE**	
Independent Variables	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***	***Coef.***	***P***
CRED	-0.3095	0.123	-0.3397	0.089	0.0256	0.889	-0.0184	0.921	0.0864	0.641	0.0798	0.663
OWNR	0.0992	0.165	0.0965	0.176	0.0146	0.81	0.0108	0.86	0.0906	0.136	0.0902	0.138
SIZE	0.0185	0.28	0.0200	0.245	0.0057	0.722	0.0084	0.604	0.0266	0.099	0.0311	0.027**
OPRC	-0.1599	0.015**	-0.1535	0.019*	-0.2087	0.001**	-0.2019	0.001**	-0.0381	0.538	-0.0476	0.426
INFL	-2.5092	0.052	-0.9599	0.07	-3.9280	0.001**	-1.4795	0.004**	1.0895	0.383		
GDPPC	-0.1386	0.079	-0.1435	0.07	-0.2161	0.004**	-0.2254	0.003**	0.0951	0.21		
MKTC	21.373	0.188			33.756	0.03**			-6.5370	0.677		
CONS	0.5773	0.604	1.7787	0.006	0.7974	0.454	2.6978	0	-0.3034	0.778	0.0735	0.803
Prob > chi2	0.0041		0.0473		0.0005		0.0021		0.1118		0.0473	
Dependent Variables	*(1) CRS TE*		*(2) CRS TE*		*(3) VRS TE*		*(4) VRS TE*		*(5) SE*		*(6) SE*	
Hausman Test	0.4065		0.2791		0.024		0.0011		0.9998		0.8410	

*This table summarizes the result of the Random-Effect OLS Regression using CRS, VRS and SE as response variables. Two different models are regressed on each response variable in order to deal with multicollinearity. The symbol "**" represents statistical significance at 5% level. Model 5 is insignificant at 5% level so we ignore it in the analysis. OPRC negatively affects CRS TE while OPRC, INFL and GDPPC negatively influence VRS TE. MKTC also influences VRS TE but in a positive direction. Of all the explanatory variables, only SIZE relates to SE and that relation is positive*

Table 8: Fixed-Effects OLS Regression Analysis – Determinants of Bank Efficiency

	(3) VRS TE		**(4) VRS TE**	
	Coef.	*P*	*Coef.*	*P*
CRED	-0.0094	0.967	-0.0764	0.741
OWNR	-	-	-	-
SIZE	-0.0142	0.432	-0.0123	0.505
OPRC	-0.1079	0.105	-0.095	0.16
INFL	-3.8961	0.001**	-1.5332	0.003**
GDPPC	-0.1528	0.044**	-0.1588	0.041**
MKTC	32.635	0.031**	-	-
_cons	0.7472	0.469	2.5904	0
Prob > F	0.0096		0.0324	

*This table shows the result of the Fixed-Effect OLS Regression using on one hand VRS as response variable and on the other hand CRED, OWNR, SIZE, OPRC, INFL, DPPC and MKTC as explanatory variables. The symbol "**" represents statistical significance at 5% level. It shows that INFL, GDPPC and MKTC influence VRS TE. However, whereas the relations of INFL and GDPPC with VRS TE are negative, that of MKTC with VRS is positive.*

CONCLUDING COMMENT

This thesis mainly examined, the TE of banks in Ghana and its determinants. More precisely, it looked at the variation of TE by bank size and whether big banks enjoyed economy of scale as an advantage over small banks. Using balanced panel data of a sample of twenty-one (21) banks from 2009 to 2013, the various analyses prove that at most eight (8) banks were technically efficiency according to the production approach of DEA – BCC model. By implication, a minimum of 61.9% of banks in the country were inefficient. The analyses further indicate that the average bank in the country was 63.89% to 80.48% as efficient as the "best" banks. This means that, given the same level of output, the average bank utilized 19.52% to 36.11% more input than the technically efficient bank. Largest banks had the highest number of technically efficient banks, followed by large banks. Even where the number of technically efficient banks was the same among largest banks and the other category of banks (the case of intermediation approach, DEA – CCR model), largest banks had the highest relative TE score. There is therefore ample evidence to support the assertion that on the average, TE varied directly in proportion to bank size within the two upper quartiles. It was also evident that big banks did not benefit from economy of scales as an edge over small banks.

All the regression models show that INFL, GDPPC and MKTC influence TE substantially. Whereas the influence of INFL and GDPPC are negative, that of MKTC is positive. This signifies that increases in INFL and GDPPC lead to a decrease in efficiency whilst an increase in MKTC leads to improvement in efficiency. With regards to company-specific variables, both Tobit regression and random-effect OLS regression points at OPRC as having a negative impact on TE; implying that increases in OPRC hinder TE. According to random-effect OLS, SIZE negatively influences SE whilst Tobit regression indicates that CRED has a negative impact on TE. By inference, SE tends to suffer as banks get bigger and increases in CRED cause TE to decline. None of the regression models establishes any substantial association between OWNR and efficiency.

RECOMMENDATIONS

Bank managers should institute measures that will improve efficiency of operating cost. This involves for instance, the maintenance of an optimum number and structure of staff and the minimization of loan impairment. We can achieve the latter through prudent and efficient credit risk management, which includes rigorous analyses of potential borrowers, regular review of interest and capital repayment schedule for existing borrowers, collateralization of lending and the use of credit limits. On the other hand, government should create a conducive economic environment to enable banks to thrive. For example, because inflation relates negatively to efficiency, there is the need to keep it at a barest minimum. Our finding supports the assertion of Khan, et al. (2001) who stated that above the threshold of 3% to 6% range, inflation significantly impedes growth of the financial sector. Finally, a negative association between GDPPC and efficiency suggest that and increase in GDPPC adversely affects efficiency. Hassan and Sanchez (2007) admit the difficulty in predicting the direction of efficiency when an economy experience growth. They added that efficiency could respond to growth tardily. We therefore recommend that government pursue policies that ensure economic growth due to its immense benefit.

The direct variation of efficient in proportion to bank size in the two upper quartiles is of relevance to existing and prospective investors in the banking sector. To existing investors, it is recommended that they all yearn to augment the size of their bank to enable them correspond to that of those in the two upper quartiles (if this is not currently the case) since banks in these categories were more efficient. We also caution prospective investors to target big banks if they are risk averse. Bank customers must equally be mindful of the predominance of inefficiency in the sector and contribute towards its reduction by being trustworthy and credit worthy. Their credit worthiness may lower non-performing assets and bad debts. It may further lead to lower lending risk and possibly lower lending rate. Trustworthy customers would contribute to efficiency by providing accurate and reliable information about themselves and their businesses. This would reduce the intensity and cost of due diligence conducted by banks before granting loans. Ultimately, both credit worthiness and trustworthiness of customers may result in a reduction in operating cost and an increase in TE.

LIMITATIONS

The limitation of this work is the reliance on only one estimation techniques (i.e. DEA) to compute the efficiencies since this may produce misleading results. Combination of models like SFA, FDH may produce a more realistic result. Secondly, the use of balance panel for the analysis resulted in the elimination of banks whose operation did not span entirely over the period under consideration. Their involvement in the sample could have yielded different results. Finally, the number of variables that influences TE exceeds those considered in this paper and therefore our findings and recommendations cannot be exhaustive.

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ASSET PRICE VOLATILITY AND EFFICIENT DISCRIMINATION IN CREDIT MARKET EQUILIBRIUM

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ABSTRACT

Significant variation in the terms and volume of lending across classes of borrowers distinguished only by qualities independent of credit risk is often interpreted as evidence of inefficient or inequitable discrimination in credit markets. Increasing accuracy in the measure of credit risk renders common theories of lending discrimination and credit rationing based on lender preferences or asymmetric information increasingly implausible. We consider a traditional model of lending with complete markets, in which equilibria may exhibit disparate loan terms or access to credit across such classes of borrowers, despite common knowledge about the parameters describing the market. Rather than evidence of inefficient equilibria owing to discrimination, however, such equilibria can arise solely from the influence of asset price volatility on participants strategically exercising the options embedded in standard debt contracts. Extending substantially different loan terms or even rationing credit to different classes of borrowers can be a rational response by value-maximizing lenders when borrower classes are correlated with the degree of price volatility exhibited by the otherwise similar assets being financed by members of each of these classes. We discuss our results in the context of actual credit markets.

JEL: G13, G15, O16

KEYWORDS: Credit Rationing, Debt Contracts, Discrimination, Stochastic Differential Game

INTRODUCTION

Observations that some borrowers encounter significantly different loan terms than other Borrowers, from whom they appear indistinguishable with respect to credit risk, are common in many economies. Similar observations of periodic credit rationing to such borrowers are also frequently alleged. If true, these events pose a significant anomaly for the efficiency of credit markets. Two types of explanations for these observations are traditionally offered. The first, popular among policymakers and the public, asserts that lender preferences over classes of borrowers distinguished by observable qualities or traits which are independent of credit risk. Assuming lenders possess market power, these preferences could lead to systematic discrimination between these classes through the offer of different terms and magnitudes of credit. The second type of explanation involves an asymmetric dispersion of information across lenders and borrowers. Both adverse selection and moral hazard can produce different loan terms and credit rationing across distinct groups of borrowers who differ by unobservable probabilities of default. These models, however, cannot explain patterns of lending discrimination between borrowers based on observable qualities that are independent of credit risk.

We offer a novel explanation of discrimination and rationing in credit markets that is fundamentally different from these. We demonstrate that allocations of credit consistent with observations of discrimination can arise in efficient credit markets, in contrast to the reliance on exogenous sources of inefficiency, such as preferences and market power or asymmetric information. Our model features a representative credit market embedded in a continuous-time economy exhibiting complete markets and arbitrage-free valuation. We augment this classical environment through consideration of the strategies chosen by lenders and borrowers bargaining over the terms of a standard secured loan. These strategies include the respective choices of the parties to exercise the options embedded in such a loan. The strategic timing of their exercise depends, in part, on the properties of the asset serving as collateral. Optimality in the respective selection of strategies yields a unique non-cooperative equilibrium in which contrasting loan terms are offered to different classes of borrowers. Each such class is distinguished only by an observable trait displayed by its members while these members appear identical by standard measures of credit risk. Evidence of systematic discrimination, commonly interpreted as inefficient or inequitable, is entirely based on these two criteria. Under the condition that lenders perceive a correlation between a given class of borrowers and the properties of the representative asset securing their loans, however, we show that endogenous discrimination in the allocating credit across different classes of borrowers will be observed in market equilibrium. Unlike the inefficient equilibria exhibited in earlier approaches involving lender preferences or asymmetric information, however, our equilibrium arises in an economy with complete markets and is efficient in its allocation of credit risk. The paper is organized as follows. We review the relevant previous research on equilibrium credit discrimination in the next section. Our model is described in the succeeding section, followed by selected remarks. Concluding remarks appear in the final section.

LITERATURE REVIEW

The significance of this paper is in its resolution of the anomaly of discriminatory loan terms and credit rationing within a market satisfying the classical assumptions of finance theory. This section compares and contrasts our analysis to previous explanations of lending discrimination and credit rationing. A traditional approach to explaining idiosyncratic lending patterns across different borrowers was pioneered by Hodgman (1960) and Freimer and Gordon (1963) , and refined in subsequent models by Jaffee and Modigliani (1969), Barro (1976), Baltensperger (1976), and others. Differential access to credit arose from an exogenous inability of lenders to price default risk across individual borrowers. While able to predict lending patterns, which, for a suitable choice of parameters, could approximate the empirical patterns commonly, interpreted as demographic discrimination, these models could not generate endogenous credit market equilibria exhibiting discrimination based on borrower characteristics.

Stiglitz and Weiss (1981) analyzed the endogenous decisions of lenders who, owing to the assumed presence of either adverse selection or moral hazard, could price risk to individual borrowers only on the average characteristics commonly exhibited by a group of borrowers, despite differences in credit risk among individual members of such groups. Second-best debt contracts incorporating screening resulted in lending terms, which differed across groups, as well as circumstances in which a group could experience the rationing of credit. More recent research, including Hart and Moore (1998), Gorton and Kahn (2000), Martin (2008), Saavedra (2015) and others, have examined the influence of monitoring costs and renegotiation in the design of contracts in credit markets with adverse selection or moral hazard. Models based on asymmetric information are unsuitable, however, in explaining discrimination owing to observable traits distinguishing borrowers of equal credit risk, since the equilibrium allocation of credit depends entirely on contracts screening borrowers only for unobservable characteristics directly germane to individual credit risk. The market equilibria in such models are necessarily inefficient in allocating risk. Our model and results are distinguished from earlier quantitative as well as anecdotal research through both endogenously deriving lending discrimination using only the traditional assumptions in finance and in generating equilibria on this basis, which are efficient in the allocation of credit risk.

METHODOLOGY

We model the negotiations of a loan contract by a representative borrower and lender with a stochastic differential game. The optimal strategies of each player are determined through the valuation functions of each party about their respective claims on the asset serving to secure the loan. Since the economy is assumed to have complete markets, we apply the standard arbitrage-free valuation method to derive a pair of linked partial differential equations and corresponding boundary conditions, which must be satisfied by the values of these contingent claims. Defining the state space of the game to be the support of all asset values and dates relevant to the decisions of the borrower and lender, we employ recursive methods to endogenously derive, through these differential equations, those subsets of this state space which represent the strategy spaces of each of the parties. A numerical solution for the game is obtained by finding those sequential paths in the state space which represent the "best-reply" strategies of the parties and which, in turn, determine market equilibrium. Each choice of parameter values for the market yields a distinct and unique equilibrium. The comparison of the properties of interest exhibited by alternative equilibria are compared and through this comparison we characterize lending discrimination in the form of loan terms and the volume of credit exchanged between lenders and different classes of borrowers.

Model

Consider a representative credit market in which participants trade, over dates t a single risky asset of value a. Participants also have access to a riskless asset with maturity T and return r. The market operates in a continuous time economy, which satisfies the assumptions of the classical asset valuation environment. In particular, this economy has a complete filtered probability space $[\Omega, \Im, P]$, where Ω represents the space of events in this economy, $\Im$ represents the corresponding filtration, a set of sequential sigma algebras $\{ \Im_t \}_{0\leq t \leq T}$ representing information available to traders at time t, and P is the actuarial probability measure over asset value and defined on Ω . All market participants observe each $\Im_t$ and base all decisions at each date t, $0 \leq t \leq T$, on this observation. All such observations are also assumed common knowledge among participants. All diffusions describing asset and contract values in the economy are also assumed to be well defined and adapted in this space. Asset markets are complete with respect to the source of risk arising in our representive credit market. It exhibits, consequently, arbitrage free valuation.

Debt contracts in this market correspond to a standard nonrecourse form under which a lender advances a unit of credit to a borrower at date 0 for which the borrower is obligated to remit both continuous coupon payments at a constant rate c until maturity T and a terminal balance $C(T)$, where $C(t)$ represents the unpaid loan balance at all dates $t \in [0,T]$. The rate premium γ paid by the borrower is determined by the constant coupon rate specified in the contract, relative to the riskless rate r. The asset being financed serves as the unique collateral securing the loan. Its value at date t, $a(t)$, evolves according to the diffusion

$$a(t) = \alpha(a,t)dt + \sigma a(t)dz(t) \quad (1)$$

where $z(t)$ is a standard Brownian motion, $\alpha(a,t)$ is the expected drift at all t and σ is a constant volatility parameter. Since these features are exogenous and cannot be influenced by either party to the loan, neither moral hazard nor adverse selection are present in this market. The borrower receives a continuous flow of value $\pi(a,t)$, representing profits net of depreciation costs, while he services the loan. Foreclosure, following a failure by the borrower to pay c at any date t, results in sale of the asset. The lender receives $max[a(t) - b(a,t), 0]$ from this sale, where $b(a,t)$ is the liquidation cost incurred by the lender, and the borrower receives any residual funds exceeding the unpaid balance $C(t)$. Bargaining over the terms of the loan contract is represented by the choice of intertemporal strategies by borrower and lender. Their respective strategy spaces include choice of the timing of any exercise of the range of options normally embedded in the loan covenants. The principal element of the space of the lender is, in the current paper, his specification of the initial amount of credit advanced, *C(0)*, and, contingent on default, the date of foreclosure. This space could, in a more complex version of the model, be augmented to include several additional options, including the right to call the loan. The principal strategic elements chosen by the borrower are the timing of both his exercise of the option to default or to prepay the existing loan balance. The strategy of the lender and of the borrower are each selected to maximize, subject to the strategy of his counteparty, the value of his contingent claim on the asset collateralizing the loan.

Denote by *L(a,t)* the values to the lender of payments received from the borrower and from his option to foreclose and denote by *B(a,t)* the value of the borrower's position in the contract, also inclusive of his options to default or prepay. Consistent with the traditional features of our economy, application of standard arbitrage pricing methods yield the respective functions *L(a,t)* and *B(a,t),* generating these values for all possible combinations *(a,t).* Solutions for these value functions, under each choice of parameters, represent the respective values of the debt and equity claims on the asset in a perfect Markovian equilibrium. Solutions to these functions satisfy a pair of partial differential equations, linked by the best-reply strategies selected by each party and by the respective boundary conditions for each equation. This pair of equations is

$$rL = \left(\frac{1}{2}\right)(a\sigma)^2 L_{aa} + (ra - \pi)\mathrm{L_a} + c + L_{\mathrm{t}}\,, \tag{2}$$

$$rB = \left(\frac{1}{2}\right)(a\sigma)^2\,(a\sigma)^2 B_{aa} + (ra - \pi)B_a + c + B_t \Big), \tag{3}$$

with the corresponding boundary conditions

$$L(\breve{a}\,,t) = max\{0, \breve{a} - \eta(\breve{a}\,,t)\}, \tag{4}$$

$$\pi(\breve{a}\,,t) - c + E_t B^*(\breve{a}\,,t) = 0, \tag{5}$$

and

$$L^*(\hat{a}\,,t) = C(t), \tag{6}$$

$$\pi(\hat{a},t) - c + E_t B^*(\hat{a}\,,t) = 0, \tag{7}$$

The term $E_t(\bullet)$ is the expectations operator under the unique equivalent martingale measure induced by our assumption of complete markets and $B^*(a,t) = (e^r dt)B(a+da, t+dt)$ and $L^*(a,t) = (e^r dt)L(a+da, t+dt)$ are the respective risk-adjusted values of the claims of the borrower and lender, discounted at the riskless interest rate. Denoting by $\bar{\bar{B}}(a,t)$ and $\bar{\bar{L}}(a,t)$ the respective values of the parties' claims if the loan terminates through the exercise of an option by either party, the terms $\breve{a}$ and $\hat{a}$ are the respective asset values triggering default and prepayment b the borrower. At these values the functions *B(a,t)* and *L(a,t)* satisfy the value-matching and smooth-pasting criteria. The value-matching condition requires the borrower's value function to be continuous at the respective asset value inducing him to default at date *t*, $\breve{a}$, or to prepay at date *t*, $\hat{a}$, as defined respectively by

$$B(\breve{a},t) = \bar{\bar{B}}(\breve{a},t), \tag{8}$$

$$B(\hat{a},t) = \bar{\bar{B}}(\hat{a},t), \tag{9}$$

while the smooth-pasting condition requires the first derivatives of $B(\bullet)$ and $\bar{\bar{B}}(\bullet)$ to be continuous at these same points. The lender's value function is required to satisfy analogous criteria at those distinct points where he would exercise his option to foreclose or, if the game allows this option, to call.

Numerical Solution Procedure

Since the finite maturity of the loan precludes an analytical solution, we characterize market equilibrium through numerical solutions for the valuation equations and boundary conditions (2)−(7). We use a recursive finite difference procedure to obtain these solutions. This requires representation of the respective strategies of the lender and borrower in terms of subsets of the underlying state space of our game. This state space is defined by all specific pairs of exogenous values *a* of the asset and corresponding dates *t* relevant to the respective strategy choices by the lender and borrower. If the set of all asset values is denoted by $\mathcal{A}$ and the compact set of all dates relevant to the loan contract by $\mathcal{T}$, then the state space of the game is $\mathcal{A} \times \mathcal{T}$, which is the support of the continuum of all possible states *(a,t)*. It is within $\mathcal{A} \times \mathcal{T}$ that the strategies chosen by the lender and borrower are nested.

Any strategy chosen by either of these parties, consequently, can, for the purposes of a numerical solution for equilibrium, be represented as regions (subsets) of $\mathcal{A} \times \mathcal{T}$. These sets comprising each feasible strategy of the lender and borrower are defined by the property that any realization of the state *(a,t)* within them triggers the exercise of an option by one of these respective parties. One element of every strategy of the lender, for example, consists of his choice of the initial balance $C(0) \in \mathcal{A}$ at the origination of the loan. A second element is his choice of a date, contingent on default by the borrower, at which he chooses to foreclose. Since, in the current model, the lender will always choose to foreclose immediately at default, the lender's strategy space can be entirely represented by the closed set $\mathcal{A} \times \{0\}$. Similarly, the strategy of the borrower can be represented in terms of two closed regions of $\mathcal{A} \times \mathcal{T}$. The first is $\mathcal{D}$, which is that subset of $\mathcal{A} \times \mathcal{T}$ containing all states at which the borrower chooses to default by ceasing the coupon payment. The analogous region in which the borrower will prepay the outstanding loan balance, should his asset be sufficiently valuable, can be denoted by $\mathcal{P} \subset \mathcal{A} \times \mathcal{T}$. The strategy selected by the borrower is the union of these respective regions, $\mathcal{D} \cup \mathcal{P}$.

The endogenous derivation of numerical solutions for *L(a,t)* and *B(a,t)* proceeds by using the values of the claims of each party. Representing $\mathcal{A} \times \mathcal{T}$ as a discrete rectangular grid of *a* and *t* values, a solution for equilibrium of our game proceeds by calculating the respective values for $L(\bullet)$ and $B(\bullet)$ at each grid point (*a,t*). These values are determined using the Crank-Nicholson method of calculating the discrete approximation for the partial derivatives in equations (2) and (3). The regions comprising the strategy of each of the parties are then found by calculating, from every point at maturity $\mathcal{T}$, the respective values of the loan to each party. This calculation is repeated at $\mathcal{T}$-1 and, in general, at each prior grid point $t-1$ accessible from each of the analogous points at *t*. These values indicate whether, at each such grid point, either party could do better by exercising his options and terminating the loan. A numerical solution for the unique equilibrium of the game, given the choice of parameter values describing the exogenous features of the market, is then obtained by using these values to find the unique 'path,' for each node at $\mathcal{T}$, of *(a, t)* values through our discrete approximation of the state space $\mathcal{A} \times \mathcal{T}$ for which the subsets we derived constitute best-reply strategies for the parties. The characterization of discrimination in the market is then obtained by comparing the terms and volume of credit for the loan contract in the equilibria generated by each of our choices of alternative parameter values.

RESULTS AND DISCUSSION

Selection of alternative sets of values of the parameters determining the stochastic evolution of the asset and institutional features of the market allow us to compare the properties of the strategies and debt and equity values in the equilibrium arising for each such set. Such comparisons allow us to measure the relative difference between loan terms and the volume of credit exchanged in each equilibrium. We can, therefore, numerically assess the existence and magnitude of systematic lending discrimination in these comparisons. Since the features most frequently cited in empirical evidence of lending discrimination, such as in analyses of HMDA data in the United States (Avery, Brevoort and Canner (2007)), are comparisons of the amount of credit obtained at loan origination and selected loan terms, we focus the presentation of our results on these two properties of credit market equilibrium. We use the initial loan balance, which is also the current value of the loan to the lender $C(0)$, to measure the amount of credit and the rate premium π to represent the terms of the loan and compare these values across the equilibria corresponding to alternative sets of parameters describing the exogenous features of the credit market. The existence of a perceived correlation between classes of loan borrowers distinguished by traits unrelated to credit risk and the parametric characteristics of the diffusion process describing the value of the representative asset securing the equilibrium loan contract for each class yields different loan terms and balances offered to the members of each such class.

The exogenous features of the market we consider include the instantaneous mean $\alpha(a,t)$and proportional volatility σ exhibited over time by the collateral asset; the net flow of value $\pi(a,t)$ accruing to equity in the asset; the cost $b(a,t)$ incurred by the lender in liquidating the asset in the event of foreclosure; and any cost $f(C(t))$ incurred by the borrower should he prepay the loan. We assume, for simplicity in the presentation and discussion of the parametric cases below, that the net revenue flow from the asset $\pi(a, t)$, liquidation costs $b(a, t)$ and prepayment costs $f(C(t))$ are all independent of time and homogeneous in their arguments. These allow us to represent conveniently the initial loan balance as a percentage of the initial value of the asset. We interpret as annual the per-period values for the riskless rate and rate premium, the asset volatility and the maturity of the loan as a convenience for the reader.

We present our results as equilibrium values of the initial loan balance $C(0)$, per corresponding rate premium γ, for alternative values of a chosen exogenous variable, holding constant all other parameters at benchmark values. The benchmark values underlying the results presented below are, respectively, a riskless annualized interest rate of $r = .03$, an annualized proportional volatility (standard deviation) in the value of the collateral asset of $\sigma = .2$; an instantaneous flow of value to equity π of ten basis points, a maturity T of five periods, a one basis point flow of coupon payments c; and both liquidation $b(a,t)$ and refinancing $f(C(t))$ costs of zero. We also specify six gridpoints, or equivalently five 'periods,' over the state space $\mathcal{A} \times \mathcal{T}$ for our calculations of the numerical solutions to equations (2) – (7).

We first consider the influence of liquidation costs on the equilibrium amounts and terms of credit and show that, above a certain threshold rate premium, the lender will rationally ration credit. Table 1 illustrates this by depicting how loan balances $C(0)$ vary with successive two percentage point increases in the corresponding equilibrium rate premium γ , as the costs b of asset liquidation in foreclosure increase from a 'low value' (10%) to a 'high' value (30%). Higher balances correspond to higher rates in each case, but, as expected, higher liquidation costs reduce the initial loan balance at each rate premium. Averaged over the 12-percentage point range of rates, initial balances are 86.34% and 78.37% at the respective low and high liquidation costs, which is approximately an eight-percentage point difference. Balances also increase, at a decreasing rate, as rates rise from 2% to 8%, but with a lower average increase as liquidation costs increase. Note, in particular, that, as liquidation costs become sufficiently high, rate increases beyond a threshold point (8% in this example) elicit virtually no increases in initial balances. Credit, after this threshold, is strictly rationed.

Table 1: Loan Terms: Effects of Liquidation Costs

Rate Premium	Case One (*B* = 10%)	Case Two (*B* = 30%)
□	*C (0)*	*C (0)*
0.02	71.70	57.60
0.04	82.70	65.00
0.06	89.00	68.00
0.08	93.00	70.70
0.10	95.50	71.80
0.12	97.20	71.90
0.14	98.30	71.90

Table1 illustrates the influence of liquidation costs on equilibrium lending terms. Case One shows that, for relatively low costs (b=10%) of liquidating the asset at default, the initial loan balance is 71.7% at a rate premium of 2%. This balance increases monotonically with this rate until it reaches 98.3% of the asset value at a 14% rate. Case Two illustrates, however, that higher liquidation costs reduce both the amount of credit available at each rate premium as well as rate at which that amount increases at each successively higher rate. Liquidation costs of 30% reduce the initial loan balance to only 57.6% at a two percent premium while the increases in the balance for equal rate increases steadily decline until, at a rate premium of 10%, the balance reaches only 71.90%. True credit rationing occurs after this point, with subsequent rate increases producing no increase in credit. Higher rates provide no increase in value for the lender because liquidation costs imply that the increased risk of default they induce outweighs any increase in the value of higher coupon payments.

Table 2 illustrates the influence of price volatility on credit available at each rate. When volatility is relatively 'low' (σ =15%), for each increase in rates balances increase, but at a rate that decreases sharply from 14% at a 2% premium to 1% at a 12% premium. Doubling the volatility, as in Case Two, reduces the amount of credit available at each corresponding rate at an average decline of 10.2 percentage points relative to the lower volatility. Successive rate increases accompany balances increasing, again at a decreasing rate. Note, however, that in contrast to the analogous comparison involving liquidation costs, the average successive increase in balances is greater when volatility is high than when it is low.

Table 2: Loan Terms: Effects of Asset Price Volatility

Rate Premium	Case One (σ = 15%)	Case Two (σ = 30%)
π	*C (0)*	*C (0)*
0.02	78.10	58.10
0.04	88.70	73.70
0.06	93.60	82.10
0.08	96.20	87.50
0.10	97.70	91.10
0.12	98.70	93.70
0.14	99.20	94.40

Analogously, Table2 illustrates the influence of annual volatility in asset value on equilibrium lending terms. Case One shows that, for relatively low volatility (σ =10%), the initial loan balance is 78.10% at a rate premium of 2%. Credit available to the borrower increases monotonically with this premium until it reaches 99.2% of the asset value at a 14% rate. This increase in credit, however, exhibits a sharply decreasing rate of increase, from approximately a 14% growth in initial balance as the premium rises from a value of 2% to 4%, to only 1% growth as the premium rises from a value of 12% to 14%. Higher volatility (σ = 30%), as expected, reduces the amount of credit available, at each corresponding value of the rate premium, to an average balance of 83% relative to an average of 93.4% at the lower volatility. ,

We now illustrate, in the last set of our selected results, the effects on the availability of credit from simultaneous variation in the volatility of asset value and costs of liquidating that asset at foreclosure. Table 3 measures these effects through four parametric combinations corresponding to those in Tables 1 and 2. In the first case, loan terms are shown for 'low' (b=10%) and 'high' (b=30%) liquidation costs, conditional on the 'low' value (15%) of price volatility. The second case repeats this same increase in liquidation costs, but now at the 'high' volatility value (30%).

Table 3: Loan Terms: Combined Effects of Liquidation Costs and Asset Price Volatility

	Volatility (15%)			**Volatility (30%)**	
	b = .10	**b = .30**		**b = .10**	**b = .30**
γ	**C(0)**	**C(0)**	γ	**C(0)**	**C(0)**
0.0200	0.7690	0.7050	0.0200	0.5230	0.4990
0.0400	0.8400	0.7660	0.0400	0.6610	0.5600
0.0600	0.8700	0.7910	0.0600	0.7400	0.6030
0.0800	0.8840	0.7980	0.0800	0.7890	0.6340
0.1000	0.8900	0.8040	0.1000	0.8220	0.6540
0.1200	0.8940	0.8090	0.1200	0.8440	0.6480
0.1400	0.8970	0.8130	0.1400	0.8590	0.6420

The effects on equilibrium loan terms of simultaneous variation in the volatility of asset value and costs of liquidating that asset at foreclosure are illustrated in Table 3. Loan terms, again represented by the rate premium and corresponding initial loan balance, are shown for four combinations of parameter values. In the first case, loan terms are shown for the same 'low' (b=10%) and 'high' (b=30%) liquidations costs as considered in Table 1, conditional on the annual price volatility of 15% used for Case One in Table 2. The second case shows these same loan terms for the same two values of liquidation cost, but now conditional on the annual price volatility of 30% used for Case Two in Table 2.

Three aspects of our results are particularly notable. First, the data in Tables 1-2 clearly illustrate, as expected, that loan balances at any given rate are considerably lower for borrowers when lenders incur higher liquidation costs at constant price volatility or when price volatility increases at constant costs of liquidation. Table 3 illustrates, in addition, that the adverse impact on credit caused by a given increase in liquidation costs or volatility is significantly worsened by a respective increase in volatility or liquidation costs. Consider, for example, that the approximately 6 percentage point decline in average balances caused by a given rise in liquidation costs from 10% to 30% at an annual volatility of 15% increases to an approximately 22 percentage point decline in average balances for the same increase in liquidation costs at an annual volatility of 30%, a difference of 16 percentage points. A similar difference in average balance decline caused by a doubling of volatility occurs when the costs of liquidation increase.

Second, the response of balances to small variations in loan rates also differs in these same situations. Our results in Table 3 suggest that a given increase in liquidation costs will have a significantly smaller effect on the rate balances grow, per unit increase in loan rates, when price volatility is relatively low than when it is high. When annual volatility is 15%, for example, an increase in liquidation costs from 10% to 30% corresponds to a 20 basis point decrease in the average growth of balances per unit rate increase, but the same increase in liquidation costs at an annual volatility of 30% corresponds to a 454 basis point decrease. The opposite is true, however, for the case of a given doubling of price volatility at low, relative to high, liquidation costs.

The third, and most striking, feature of our results is that not only can the 'rationing' of credit occur to different borrowers, based on the characteristics of the assets used to collateralize their respective loans, but, for possible, if extreme, parameter values in the market, increases in the rate premiums that rationed borrowers are willing to pay can actually <u>reduce</u> the loan balance they receive. A comparison of Tables 1 – 3 illustrates that, as rate premiums rise, the successive increases in balances decline at any combination of liquidation costs and price volatility. This declining responsiveness of balances can appear as incipient or actual credit rationing. However, when liquidation costs and price volatility are both high, Table 3 demonstrates that, above a threshold rate of 10%, each successive increase of two percentage points in rates actually causes credit to decline by 60 basis points.

Each of the situations respectively illustrated in Tables 1 – 3 is, arguably, empirically plausible. In actual credit markets, capital investments will be made by borrowers exhibiting observable traits that are commonly regarded as being independent of credit risk but which are correlated with the qualities of the assets borrowers, who have a common observable trait, are financing. Such investments include residential or commercial property, various types of physical capital used by entrepreneurs involved in various kinds of small business, and proprietary designs or patents associated with highly innovative and technologically advanced sectors of the economy.

The most extensively documented example of this occurs in the American market for residential mortgage loans (Ross (1996), Nickerson, Nebhut and Courchane (2000), Li, Hossain and Ross (2010) and many others.) Consider a situation in which two borrowers possessing observable differences in demographic or other traits seek financing in order to purchase a house. Assume that, on the basis of such conventional underwriting measures as an individual's financial capacity, credit history and employment status, these borrowers appear to pose similar degrees of credit risk to a lender. Borrowers with these contrasting traits, however, each tend to reside in geographically distinct neighborhoods. The average property in the neighborhood of a borrower with trait A is known, for a variety of reasons, to have higher liquidation costs per dollar, or greater price volatility or poorer information about its selling price over time, than the average property in the neighborhood of a borrower with trait B. Correlation between the demographic traits of these two types of borrowers and the respective characteristics of the properties they wish to finance will inevitably lead to a consistent difference in loan terms obtained from a given lender and each of these two types of borrowers. Observations of statistically significant differences in loan terms offered to each type would, based on public opinion as well financial regulations in a variety of countries, be interpreted as discrimination on a basis independent of credit risk. Contrary to this common interpretation, however, our model and its results prove that such systematic differences can occur in markets embodying the traditional assumptions of financial theory and be consistent with an efficient allocation of credit.

CONCLUDING REMARKS

Empirical evidence exists of discrimination in the credit markets of many economies, including those of the U.S. and E.U. This evidence largely consists of observations of significant differences in lending terms among different borrowers who, by conventional measures, appear identical in the credit risk they pose to lenders. These observations are most often interpreted as discrimination on the basis that borrowers are distinguished by contrasting demographic characteristics rather than on any objective economic basis. If this interpretation is correct, this evidence poses a serious anomaly for the traditional presumption by financial economists that the market allocation of credit is efficient.

Previous explanations of this evidence share the presumption that some form of market failure pervades lending and that, consequently, credit markets necessarily exhibit allocational inefficiencies. In stark contrast, this paper offers an explanation based on complete markets, arbitrage-free valuation and other features standard in traditional valuation models. By taking into account the strategic interaction between lenders and borrowers in negotiating loan terms, we demonstrate that the observable properties of non-cooperative equilibria in credit markets depend heavily upon price volatility and other exogenous characteristics of the asset securing a loan. Although absent from standard statistical measures of credit risk, differences in these characteristics induce both lenders and borrowers to change the conditions under which they strategically choose to exercise the options available to them in a standard loan contract. These changes, in turn, produce equilibria with significant differences in loan terms and available credit.

If lenders perceive a correlation, unaccounted for in existing credit discrimination data, between an observable but intrinsically irrelevant trait distinguishing borrowers and the properties of the representative asset securing the borrower's loan, market equilibrium will display differences in the terms and amount of credit obtained by these different borrowers. These differences will, for plausible parameter values, be observationally equivalent to those in the empirical evidence of discrimination in credit markets, but will be consistent with an efficient allocation of credit among all borrowers. Our paper also demonstrates the feasibility of applying the contingent claims method of valuation to the analysis and explanation of differences in credit available to borrowers based on factors other than credit risk. Such applications could significantly widen our consideration of factors in future research on markets for credit, and, of equal importance, offer a new basis for the design and assessment of financial regulations.

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