Macroeconomic Shocks and the Interrelations between Bank Credit and Securities Issuance

Fabrizio Casalin and Enzo Dia
Newcastle University Business School, NE1 7RU, Newcastle upon Tyne, UK
University of Milano – Bicocca,,Department of Economics,20146, Milano, ITALY.

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Macroeconomic Shocks and the Interrelations between Bank Credit and Securities Issuance

Fabrizio Casalin* and Enzo Dia†

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Abstract

This paper provides an empirical analysis of the impact of macroeconomic factors on the aggregate volume of C&I loans, and on the issuance of stocks and bonds, for the US economy. Our results suggest that the three aggregates are substantially interrelated, forward looking, and influenced by business and monetary policy indicators. We then study the levels of interrelation among the three markets by making use of conditional correlation coefficients. These last depend on business and monetary policy indicators as well as on measures of financial development: Lower levels of both debt and stock market capitalization, as well as macroeconomic and financial shocks such as the 1981 and 1991 economic recessions, the 1987 stock market crash as well as the burst of the dot-com bubble all weaken the co-movement among the three markets, providing substantial evidence against the hypothesis of financial contagion.

Keywords: Security Issuance, Credit Market, Secondary Market Prices, Interrelation, Contagion.

JEL classification: C32, G01, G12.

*Newcastle University Business School, NE1 4SE, Newcastle upon Tyne, United Kingdom. E-mail address: fabrizio.casalin@newcastle.ac.uk. Tel.: +44-0191-202-1595, Fax:+44-0191-222-6838.

†Università degli Studi di Milano Bicocca, Department of Economics, P.zza dell'Ateneo Nuovo, 1, 20146, Milan, Italy. E-mail address: enzo.dia@unimib.it.

I Introduction

The global scale of episodes such as the October 1987 stock market crash and the dot-com bubble have created a growing impetus among researchers to investigate the various aspects of the interrelation among financial markets, and to identify the channels through which such shocks spread from one market to the others. While there is a large body of literature which investigates such linkages among secondary markets, the issue of studying the levels of interrelation between primary markets of securities has remained largely unexplored.[1] The investigation of how shocks are transmitted across primary markets is important in many respects. For instance, in addition to affecting asset prices, substantial changes in the risk generated by financial markets, political climate as well as the occurrence of economic slowdowns can also affect the willingness of investors to extend funds to corporations. It is possible that movements of prices in secondary markets is only part of a more complicated dynamic in which shocks spread also through primary markets via correlated withdrawals of funds. Moreover, primary markets for securities constitute important sources of finance for the economy. The understanding of their dynamics as well as of the extent to which they are interrelated is important to asses the resilience of the economic system against macroeconomic and financial shocks that might cause sudden dry up of liquidity. We fill this gap by providing an empirical analysis of the different dynamics of the aggregate volumes of share (STOCKS) and bond (BONDS) issuance for non-financial firms and levels of commercial and industrial (C&I) loans outstanding, for the US economy. We disregard long-run trends, focussing instead on the short-term dynamics, to study how macroeconomic determinants influence the equilibria of the three markets in a simultaneous equations framework. This approach is useful as, for instance, it makes it possible to investigate whether the empirical results obtained in previous studies for share IPOs still hold when jointly evaluated with other types of security placements (see e.g. Lowry 2003). The second contribution of the present study to the existing literature consists of a through analysis of the macroeconomic and financial factors as well

[1] See Pericoli and Sbracia (2003) for an survey of the literature.

as external shocks that can explain the changing correlation patterns across markets and over time. This analysis is conducted into two different steps. We initially carry out a joint estimation of the conditional correlation coefficients among the three markets by using multivariate GARCH models. We then regress the estimated conditional correlation against macroeconomic and financial factors that can explain the time varying nature of the linkages among the three markets. For instance, the identification of significantly increased correlations among the volumes of the three aggregates subsequent to the occurrence of external shocks such as the October 1987 stock market crash could be regarded as evidence of contagion.

Several important findings stem from our analysis. First, we find strong evidence that the primary markets for shares and bonds, and C&I loans are interrelated and that business conditions indicators explain a sizeable proportion of the variability of the volumes of securities traded in the three markets.[2] For instance, both the issuance of corporate bonds and C&I loans react to changes in the rate of economic activity and monetary policy stance, but in opposite ways: bond issuance rises when short-term interest rates and the industrial production decline, whereas the opposite result holds for C&I loans. Thus, firms react to worsening market conditions initially by issuing securities, and by recurring to bank lending, for example commitment loans, only when market conditions become more favorable. These results suggest that large industrial corporations that have access to bond markets may substitute bank lending with market funds. As a consequence, the development of an efficient market for corporate bonds is important not only to reduce the market power of banks, but also to increase the resilience of the economic system against macroeconomic shocks. Second, the estimated conditional correlations among the three markets are statistically significant and exhibit significant variation throughout the sample with substantially different dynamics. On the one hand, the conditional correlation between BONDS and C&I loans is negative for the entire sample, providing further support to the hypothesis that the two aggregates are substitutes. On the other hand, the

[2]The results we obtain for primary placements of shares are in line with the existing literature. See, e.g., Lowry (2003) and Ivanov and Lewis (2008).

correlation between BONDS and STOCKS is positive, indicating strong co-movement between the two aggregates over time. The conditional correlation between C&I loans and STOCKS displays values that fluctuate from positive to negative over time, suggesting the existence of a weak linkage between the two securities. Third, we identify a number of macroeconomic and financial factors that can explain the time-varying levels of interrelation among the three aggregates. Finally, we find that macroeconomic and financial shocks such as the 1981 and 1991 economic recessions, the 1987 stock market crash as well as the burst of the dot-com bubble consistently weaken the linkages among the three markets, providing substantial evidence against the hypothesis of financial contagion.

The remainder of the paper is organized as follows. Section II discusses the relevant literature. Section III describes the dataset. Section IV introduces the empirical methodologies used. Sections V and VI discuss the empirical results obtained while Section VII concludes.

II The literature

Our empirical study draws on a number of different strands of literature. A large body of empirical literature has investigated the linkages between secondary markets of shares and bonds for developed and emerging economies (see, e.g., Connolly et al. 2007 and Syllignakis and Kouretas 2011). There is now convincing evidence that the processes of financial markets co-movement are actually dynamic, and a number of different empirical techniques have been proposed to account for this feature. Earlier studies have analyzed the levels of interrelation among secondary markets by means of unconditional cross-market correlations on prices or returns, as well as multivariate vector auto-regression and co-integration techniques (see, e.g., Corsetti et al. 2005 and Voronkova 2004). However, as noted by Kim et al. (2005), the long-run stable equilibrium relationships conjectured by the above methodologies are not suitable for modelling processes which seem to be dynamic and time varying in nature. For this reason, in the last decade scholars have made a widespread use of a variety of univariate and multivariate GARCH models to investigate phenomena such as interrelation and spillover effects among returns on stock and bond markets. These studies make use of conditional correlation coefficients to estimate the extent of co-movement among prices on secondary markets, and show that such interrelations are often strong, time-varying and dependant on macroeconomic and financial factors such as the rate of economic activity as well as monetary policy stances (see, e.g., Wang and Moore 2008).

In the above literature there is a considerable amount of ambiguity concerning the definitions of interdependence (or co-movement) and contagion. Only recently researchers have introduced a clearer distinction between the terms by referring to normal interdependence as crises that propagate because of fundamental real and financial cross-market links, and to contagion as shocks that yield a discontinuity in the data generating process of prices and quantities. While a number of different definitions of contagion have been put forward, we choose to adopt that proposed by Pericoli and Sbracia (2003) according to which contagion is a significant increase in co-movements of prices and quantities across markets conditional on a shock occurring in one market.[3]

A second strand of research important for our study is that which aims to explain the underlying causes of the dramatic swings in the volume of primary placements of shares and bonds observed in developed financial markets. The main empirical results suggest that macroeconomic factors are important driving forces for external finance, so that demand side factors should not be disregarded. These factors, in fact, could directly influence stock and bond market prices, and indirectly the amount raised as well as the timing of the issuance. In line with this hypothesis, researchers have put forward a number of different explanations of the cyclical nature of IPOs based on changing business conditions (e.g. Pástor and Veronesi 2005), investor sentiments (e.g. Dorn 2009), and asymmetric information between owners and outside investors (e.g. Dittmar and Thakor 2007). These

[3]On the other hand, any continued high degree of market correlation implies only interdependence. Therefore, contagion must involve a dynamic increase in correlations in the aftermath of crisis episodes. See Pericoli and Sbracia (2003, p.575).

studies differ in terms of dataset used and results obtained. In terms of dataset, empirical analysis are carried out by using annual and quarterly aggregate data on IPO volumes (e.g. Lowry 2003 and Pástor and Veronesi 2005) whereas data at monthly frequency consist of the number of transactions (see e.g. Ivanov and Lewis 2008). In terms of results obtained, all these studies find supporting empirical evidence for the hypothesis that capital demand lies behind fluctuations in aggregate IPO volumes, while they obtain mixed results regarding the asymmetric information and investor sentiments hypotheses. For instance, Lowry (2003) considers closed-end fund discounts and dispersions of abnormal returns around earning announcements as proxies of investor sentiments and asymmetric information, and shows that only the former has explanatory power on the volume of share IPOs. Other studies such as Hovakimian et al. (2001) find that high past stock returns and market-to-book ratios are associated with larger issues of common equity as well as long-term debt (see also Ritter and Welch 2002 and Pagano et al. 1998).

Finally, the impact of macroeconomic variables on C&I loans has been documented by a number of different studies, several of which have a specific focus on the credit channel for the transmission of monetary policy. More specifically, Gertler and Gilchrist (1993) and, more recently, Den Haan et al. (2007), find that, in the case of large banks, C&I loans issuance rises following a monetary tightening. On the contrary, Kashyap et al. (1993) document a marginally significant negative impact of short-term interest rates on C&I loans, matched by an opposite and similar impact in magnitude on commercial papers.

III Dataset

The dataset gathers monthly aggregate data for the volumes of primary placements of shares (STOCKS) and bonds (BONDS) of non-financial corporations, and the volume of all commercial and industrial loans at all commercial banks (C&I) for the US economy.[4] The dataset includes also series for the Industrial Production Index, the Composite

[4] The volumes of C&I loans represent 19% of total aggregate loans at the end of 2009.

Index of Leading Indicators, yields on three-month T-Bills, a yield spread between ten- and three-year government bonds, returns on the S&P500 and Barclays Corporate Bonds Index (BCBI).[5] The period under analysis spans from January 1973 to June 2007 for all the series. We choose not to include the period after June 2007 as characterized by severe financial distress and abnormal volatility in all the above three aggregates.[6]

The upper panel of Table 1 reports the Augmented Dickey-Fuller (ADF) as well as the Phillips-Perron (PP) tests for the null of unit-root in the levels of the series taken in log. From a visual inspection, the series do not present clear mean reverting patterns. However, when the two tests are applied to STOCKS and BONDS the null of non stationarity is soundly rejected. This finding is quite surprising and it should be taken with caution as it is a well-known result that the distribution of standard unit-root tests critically depend on the assumption that the underlying stochastic process is purely autoregressive. In fact, Schwert (1989) shows that when the underlying process contains a moving average component the distribution of unit root test statistics can be far different from those reported by Dickey and Fuller.[7] To investigate whether this is the case we fit ARIMA(0,1,1) models to the placements of stocks and bonds and find that the moving average parameters θ are equal, respectively, to 0.569 and 0.694. Given these values of the parameters θ, the corrected critical values for the ADF and PP statistics are -8.902 and -12.24.[8] By comparing the ADF and PP statistics reported in Table 1 with these critical values we conclude that the null of unit root cannot be rejected at the 5% level. The ADF and PP tests are then applied to log-levels of C&I loans and clearly suggest that the null of unit-root cannot be rejected at standard significance levels.[9]

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[5]These series are obtained from the Federal Reserve Bulletin, Federal Reserve Bank of St Louis, OECD, and Datastream. All series are deflated by CPI.

[6]The extension beyond June 2007 would make problematic the empirical estimation of our econometric models as even specifications which include a large number of parameters and lags would fail to account for the severe degree of serial correlation and heteroscedasticity induced in raw and standardized residuals.

[7]More specifically, the author shows that if the process is an ARIMA(0,1,1) with a large moving average parameter θ most of the tests depart from the distribution calculated by Dickey and Fuller and that the bias becomes more and more severe as the parameter θ becomes greater than 0. By conducting Monte Carlo experiments the same author provides corrected critical values which accounts for the presence of the moving average component.

[8]These critical values are calculated by interpolating the figures reported in Schwert (1989) pag. 89.

[9]The results obtained for the three series are robust for different specifications of the two unit-root tests.

The middle and lower panels of Table 1 report also some descriptive statistics for the three series taken in log first differences. The sample moments for the C&I loans indicate empirical distributions with fat tails relative to the normal distribution as the null of zero excess kurtosis is soundly rejected at standard significance levels. The Ljung-Box Q statistics applied to the raw series suggest the presence of strong serial correlation in all the three aggregates. The Q statistics are then applied to detect serial correlation in the squared series. In this case they consistently reject the null for the series STOCKS and C&I suggesting the presence of nonlinear dependence, possibly due to changing conditional volatility over time. This last result, however, does not hold for BONDS, suggesting that this last series is homoscedastic. We then compute the Q statistics for leads and lags of raw and squared series in order to test for the presence of lead/lag serial correlation in first and second moments. The result of these tests (not reported to save space) indicates the presence of weak interactions among the first moments of the series. Moreover, the interaction substantially vanishes among the second moments. This last result suggests the absence of volatility spill-overs among the aggregates considered.

Table 1: Summary Statistics for the Monthly Volumes of Placements of Stocks and Bonds, and C&I loans.

	STOCKS		**BONDS**		**C&I**	
	statf	**p-value**	**stat**	**p-value**	**stat**	**p-value**
ADF†	-4.668	(0.001)	-5.228	(0.001)	-3.006	(0.132)
PP‡	-8.227	(0.000)	-11.19	(0.000)	-2.187	(0.495)
Mean	0.004	-	0.003	-	0.002	-
Std Dev	0.436	-	0.429	-	0.007	-
Skewness	-0.368	(0.001)	0.047	(0.682)	0.168	(0.148)
Kurtosis	0.355	(0.125)	0.180	(0.439)	0.656*	(0.005)
Obs	449	-	449	-	449	-
Q(4)	65.82	(0.000)	72.33	(0.000)	427.3	(0.000)
Q(8)	67.12	(0.000)	82.91	(0.000)	646.3	(0.000)
Q^2(4)	18.75	(0.000)	8.846	(0.065)	46.68	(0.000)
Q^2(8)	28.58	(0.000)	9.133	(0.331)	48.87	(0.000)

Notes: Sample period 1973:01 - 2007:06. † and ‡ indicate Augmented Dickey-Fuller and Phillips-Perron unit-root tests applied to the log of the series in levels. Q(n) and Q^2(n) are Ljung-Box statistic for serial correlation in raw and squared log first differences up to lag n.

IV The Model

The empirical analysis is conducted by following two different approaches. We initially carry out a simultaneous estimation of the equilibrium quantities of STOCKS, BONDS and C&I loans in order to isolate the different impacts of our chosen macroeconomic determinants on the three markets. We then make use of a multivariate GARCH model to analyze the stochastic properties of the second moments of the three equilibrium quantities and to estimate their co-movement over time.

We do not aim to explore a detailed structural model of the banking industry or of the primary markets for securities. We rather focus on reduced form equilibria by regressing the equilibrium quantities on variables different from the respective market prices, capturing the cyclical behavior of aggregate demand and interest rates. Similar reduced form equilibria can be obtained from standard theoretical models. More specifically, when analyzing loan aggregates, the dynamic structure we assume can be obtained from standard dynamic models of banking under the assumptions that interest rates and aggregated demand can be described as random walk processes, so that the expectations of future values are captured by the current values of the same variables (e.g. Cosimano 1988 and Chami and Cosimano 2010). The equations describing the market for securities are standard in the literature on primary placements (see, e.g., Lowry 2003 and Pástor and Veronesi 2005). This literature shows that business conditions are important determinants for the issuance of shares and debt.[10] In the light of this evidence, we consider an expanded set of proxies for business conditions which includes indicators for the expected cost of capital and profitability. Private firms, in fact, respond to changing market conditions by optimally choosing to go public when the expected cost of capital is low. Moreover, time-variation in expected profitability creates periods in which firms find it desirable to raise finance so they can exercise growth options. First-differences in three-month T-Bills yields (the risk-free rate, $\mathbf{i}_t$) and realized returns on the stock ($\mathbf{R}_{S\&P,t}$) and bond ($\mathbf{R}_{BCBI,t}$)

[10] The business condition hypothesis asserts that during economic expansions, economy-wide demand for capital is higher and more firms go public.

secondary markets are the proxies used for the expected cost of capital (see, e.g., Mayfield 2004 and Helwege and Liang 2004). The indicators we use to proxy the expected profitability are the log first differences in the levels of the Composite Leading Indicators ($\mathbf{CLI}_t$) and Industrial Production indices ($\mathbf{IPI}_t$), as well as the yield spread $\mathbf{YD}_t$ (see, e.g., Pástor and Veronesi 2005 and Ivanov and Lewis 2008). We study the aggregates in log first differences since they are non-stationary, and we deflate the data by using the CPI in order not to capture the dynamics of inflation.[11] The specification of the first empirical model we employ is as follows:

$$\Delta S_t^1 = \alpha_0^1 + \alpha_2^1 \Delta S_t^2 + \alpha_3^1 \Delta S_t^3 + \alpha_4^1 R_{S\&P,t} + \alpha_5^1 \Delta CLI_t + \alpha_6^1 R_{S\&P,t-1} + \sum_{j=1}^{3} \sum_{i=1}^{L_j^1} \gamma_i^j \Delta S_{t-i}^j + \varepsilon_t^1 \quad (1)$$

$$\Delta S_t^2 = \alpha_0^2 + \alpha_1^2 \Delta S_t^1 + \alpha_3^2 \Delta S_t^3 + \alpha_4^2 R_{BCBI,t} + \alpha_5^2 \Delta IPI_t + \alpha_6^2 \Delta i_t + \sum_{j=1}^{3} \sum_{i=1}^{L_j^2} \gamma_i^j \Delta S_{t-i}^j + \varepsilon_t^2 \quad (2)$$

$$\Delta S_t^3 = \alpha_0^3 + \alpha_1^3 \Delta S_t^1 + \alpha_2^2 \Delta S_t^2 + \alpha_4^3 \Delta IPI_t + \alpha_5^3 \Delta i_t + \alpha_6^3 YD_t + \sum_{j=1}^{3} \sum_{i=1}^{L_j^3} \gamma_i^j \Delta S_{t-i}^j + \varepsilon_t^3 \quad (3)$$

where ΔS_t^j are the log first differences in the volumes of the security j (for j=1,2,3), and L_j^1, L_j^2, L_j^3 define the number of lags and cross-lags included in each equation of the system of eqs. (1)-(3).[12]

The statistics reported in Table 1 suggest the presence of both strong serial correlation and changing conditional volatility in the three aggregates. The second model we employ aims to account for both these features of the data. More specifically, we make use of standard multivariate GARCH frameworks to model the growth rates of the different aggregates as well as their conditional covariances matrix. This type of models has been extensively used in the literature on the linkages among financial markets. In line with this literature, we assume that the mean equations follow a VAR(p) stochastic process in

[11] The empirical results for the nominal variables are very similar to those reported in Table 2, but characterized by stronger statistical significance. These results are not reported but are available from the authors.

[12] The parameter j takes values 1 for STOCKS, 2 for BONDS and 3 for C&I loans. The inclusion of lagged dependent variables is in line with Granger and Newbold (1974) and with previous studies on IPOs such as Lowry (2003).

which each equation is specified as follows:

$$\Delta S_t^1 = \mu_1 + \sum_{j=1}^{3} \sum_{p=1}^{P_j^1} \gamma_{j,p} \Delta S_{t-p}^1 + \varepsilon_t^1 \tag{4}$$

$$\Delta S_t^2 = \mu_2 + \sum_{j=1}^{3} \sum_{p=1}^{P_j^2} \gamma_{j,p} \Delta S_{t-p}^2 + \varepsilon_t^2 \tag{5}$$

$$\Delta S_t^3 = \mu_3 + \sum_{j=1}^{3} \sum_{p=1}^{P_j^3} \gamma_{j,p} \Delta S_{t-p}^3 + \varepsilon_t^3 \tag{6}$$

where the growth rates of the three aggregates depend on a constant μ_i, on their own P_i^j lags and cross lags, and on the terms $\varepsilon_{i,t}$ that capture the ”unexpected shocks” on the dependent variables. The conditional covariances matrix Σ_t is assumed to follow a standard Diagonal VECH GARCH(1,1) model which can be written as follows:

$$vech(\Sigma_t) = \mathbf{C} + \mathbf{A}\, vech(\varepsilon_{t-1}\varepsilon_{t-1}^{'}) + \mathbf{B}\, vech(\Sigma_{t-1}) \tag{7}$$

where the matrices Σ_t and $\varepsilon_t\varepsilon_t^{'}$ are symmetric of dimension (3 $\times$ 3), the vector **C** has dimension (6 $\times$ 1) and both the matrices **A** and **B** are symmetric of dimension (6 $\times$ 6).[13] In order to reduce the large number of parameters under estimation we introduce the assumption that the matrices **A** and **B** are diagonal. Thus, Eq.(7) can be rewritten, after conveniently rearranging the parameter indices, as follows:

$$\sigma_{ij}(t) = c_{ij} + a_{ij}\, \varepsilon_{i,t-1}\varepsilon_{j,t-1} + b_{ij}\, \sigma_{ij,t-1} \tag{8}$$

for (i,j)=1,2,3. Given a vector of dimension (3 $\times$ 1) which collects the series ΔS_t^j and a sample of **T** observations, maximum likelihood estimates are obtained by making use of the Broyden-Fletcher-Goldfarb-Shanno (BFGS) algorithm.

[13]Since the number of series under analysis is limited to three we choose to estimate a standard multivariate Vech GARCH model instead of more sophisticated specifications such as the DCC GARCH. The operator Vech stacks columns of the lower triangle of the matrices Σ_t and $\varepsilon_t\varepsilon_t^{'}$ in vectors of dimension (6 $\times$ 1).

V Empirical Results

V.1 System estimation

The hypotheses we want to investigate involve linear relationships among the log first differences of the aggregate volumes of primary placements of shares (**STOCKS**), corporate bonds (**BONDS**) and C&I loans (**C&I**), plus a set of pre-determined explanatory variables taken from the literature on the IPOs of shares and bonds and that on banking. We follow a general-to-specific approach in which the model of eqs. (1)-(3) is initially estimated with lags and cross-lags of dependent variables as well as lagged stock and bond returns up to the sixth lag.[14] The lags not statistically significant are then removed, and the model re-estimated by means of Three Stages Least Squares (3SLS). This procedure is re-iterated until all the lagged variables included in the system are statistically significant.[15] Table 3 reports the empirical estimates carried out by using both the 3SLS and GMM estimating procedures (with White's heteroscedasticity consistent covariance matrix) as well as diagnostic tests in the lower panel.[16] The results obtained under the two estimation techniques are very similar. In both cases the residuals appear reasonably well behaved. Both the Ljiung-Box Q statistics and LM tests suggest weak presence of serial correlation in the residuals. Moreover, both the ARCH LM tests and the Q statistics applied to squared residuals show moderate presence of heteroscedasticity. Adjusted R-squared for the three regressions are reasonably high, even when omitting the autoregressive terms. The Wald F-statistics for the null that the macro factors do not exert any

[14] The order of lags for the lagged and cross-lagged dependent variables was chosen by estimating a standard VAR model for the variables ΔS_t^1, ΔS_t^2 and ΔS_t^3 and by taking the longest lag length among those suggested by the LR statistic, the Akaike, Schwarz and Hannan-Quinn information criteria.

[15] The final estimates are characterized by the following lag structure: $L_1^1 = 5$, $L_2^1 = 3$, $L_3^1 = 1$, $L_1^2 = 1$, $L_2^2 = 5$, $L_3^2 = 3$, $L_1^3 = 1$, $L_2^3 = 1$, $L_3^3 = 5$. The instrumental variables considered are lagged values of primary placements of stocks, bonds and C&I loans, current and past stock and bond market returns as well as current values of the yield spread, changes in Industrial Production, Composite Leading Indicators index and yields of three-month T-Bills. The model has been supplemented with dummy variables to account for monthly seasonality and for a number of idiosyncratic shocks such as the Stock Market Crash of October 1987 and the collapse of LTCM of 1998.

[16] We do not report the results for the lags and cross lags beyond the first, as we are not interested in the long-run responses of the variables.

explanatory power are soundly rejected at standard significance levels. For instance, the null tested for the first equation of the model is $H_0 : \alpha^1_2 = \alpha^1_3 = \alpha^1_4 = \alpha^1_5 = \alpha^1_6 = 0$; this hypothesis is soundly rejected as the χ^2-statistic calculates to 54.86 with p-value 0.000. Similar evidence is obtained for the second and third equation.

In line with the literature, we find that current and lagged values of the returns of the S&P500 as well as the CLI drive primary placements of shares (see, e.g., Ivanov and Lewis 2008). In terms of economic impact, a one standard deviation increase in the return index boosts the growth rate of the issuance of shares by 5.5% in the same period and 5.1% in the subsequent.[17] The new result we find is that an acceleration in the issuance of bonds significantly and positively influence the issuance of shares, and vice-versa. Both the contemporaneous and lagged coefficients measuring the dynamic links between the two aggregates are positive and statistically significant at the 1% level. In this case, a one standard deviation increase in the issuance of bonds accelerates the issuance of shares by 24.3% in the same period and 12.6% in the subsequent. This finding highlights the existence of a strong relationship of complementarity between the issuance of the two securities and it suggests, in line with Lowry (2003), that common factors, such as technological shocks, can drive the issuance of both. On the contrary, we find that bond issuance and changes in the levels of C&I loans have a negative contemporaneous impact on each other, whereas the long-run impact is positive in both cases. Primary placements of bonds are associated with positive returns on the bond index (BCBI) (indicating lower interest rates), and declining T-Bill rates.[18] Thus, corporate bond issuance rises following expansionary monetary policy. The Industrial Production Index is also significant, and the sign of the attached coefficient is negative, indicating that the issuance of corporate bonds is counter-cyclical. A one standard deviation increase in the Industrial Production Index is associated with an decrease in the issuance of bonds of 4.1%.

The main driving force behind the issuance of C&I loans is the short-term interest rate,

[17] Figures are expressed on a monthly basis.

[18] A one standard deviation increase in the bond index is, in fact, associated with an acceleration in the issuance of bonds equal to 3.7%.

Table 2: 3SLS and GMM Empirical Estimates of the Model of Eqs.(1)- (3).

	3SLS			**GMM**		
	STOCKS$_t$	**BONDS**$_t$	**C&I**$_t$	**STOCKS**$_t$	**BONDS**$_t$	**C&I**$_t$
STOCKS$_t$	–	0.403*** (0.072)	0.004** (0.001)	–	0.323*** (0.078)	0.003* (0.001)
BONDS$_t$	0.558*** (0.122)	–	-0.003**. (0.001)	0.499*** (0.134)	–	-0.002 (0.002)
C&I$_t$	32.85*** (8.868)	-20.65** (10.10)	–	21.76** (10.14)	-13.66 (12.78)	–
STOCKS$_{t-1}$	-0.529*** (0.049)	0.202*** (0.048)	0.002*** (0.001)	-0.550*** (0.053)	0.167*** (0.049)	0.0015* (0.001)
BONDS$_{t-1}$	0.288*** (0.086)	-0.611** (0.045)	-0.002*** (0.001)	0.256*** (0.096)	-0.620*** (0.044)	-0.001 (0.001)
C&I$_{t-1}$	-25.24*** (6.138)	22.76*** (5.050)	0.516*** (0.051)	-17.17** (7.540)	18.22** (7.120)	0.478*** (0.064)
Δ **CLI**$_t$	6.732* (3.616)	–	–	11.17** (4.700)	–	–
Δ **IPI**$_t$	–	-5.125** (2.076)	0.040 (0.038)	–	-4.460* (2.460)	0.058 (0.041)
Δ **i**$_t$	–	-0.092** (0.036)	0.002*** (0.001)	–	-0.085* (0.044)	0.020*** (0.001)
R$_{BCBI,t}$	–	1.6837** (0.732)	–	–	2.192*** (0.849)	–
R$_{S\&P,t}$	1.187*** (0.413)	–	–	1.213** (0.475)	–	–
R$_{S\&P,t-1}$	1.102*** (0.374)	–	–	1.490*** (0.520)	–	–
R^2	0.21	0.43	0.43	0.33	0.48	0.45
Q(4)	2.408 (0.492)	6.270 (0.099)	0.639 (0.887)	2.459 (0.482)	4.545 (0.208)	1.038 (0.792)
Q(8)	4.325 (0.741)	12.19 (0.094)	1.927 (0.963)	6.449 (0.488)	11.70 (0.111)	3.019 (0.883)
LM(4)	6.738 (0.874)	11.92 (0.108)	5.937 (0.919)	7.112 (0.850)	7.113 (0.850)	6.390 (0.895)
Q^2**(4)**	6.591 (0.086)	2.117 (0.548)	9.091 (0.028)	5.570 (0.143)	2.268 (0.518)	13.18 (0.004)
Q^2**(8)**	16.53 (0.021)	10.57 (0.158)	10.09 (0.183)	15.98 (0.025)	11.04 (0.136)	13.44 (0.062)
ARCH(4)	7.161 (0.127)	2.011 (0.733)	8.668 (0.069)	5.961 (0.202)	2.237 (0.692)	13.60 (0.093)

Notes: Sample period 1973:01 - 2007:06. Estimates of constant term for the three regressions not reported. GMM is estimated with White's heteroscedasticity consistent covariance matrix. Standard errors are in parenthesis. * = significant at 10%, ** = at 5% and *** = at 1%. Adjusted R^2 calculated as $1-(1-R^2)/(T-1/T-k)$. Q(n) and Q^2(n) are Ljung-Box statistics for serial correlation up to lag n in the raw and squared residuals. $\flat$ LM test for serial correlation in residuals up to lag 4. $\natural$ ARCH LM test for heteroscedasticity in residuals up to lag 4. P-values in parenthesis.

as loans outstanding grow as interest rates rise. This result is in line with the findings of Gertler and Gilchrist (1993) and Den Haan et al. (2007) that, in the case of large banks, C&I loans issuance rises following a monetary tightening. C&I loans are also positively influenced by the industrial production, indicating that the volume of loans is pro-cyclical, although the attached coefficient is not statistically significant. While bank lending declines in economic downturns, the issuance of corporate bonds grows. In a similar fashion, when monetary policy is expansionary, bond issuance grows, whereas direct lending decreases. In this case, a one standard deviation increase in short-term interest rates (equivalent to 51 basis points) is expected to accelerate the growth rate of C&I loans by 0.1% and to decrease that of the issuance of bonds by 4.6%.[19] To provide an example, the volume of C&I loans outstanding in real terms in January 2000 was 5924.4 billions of US dollars, whereas the issuance of bonds was 86.23 billions (the average bond issuance in real terms for the year 2000 was 116.9 billions). The magnitudes involved are therefore an increase of 5.92 billions in the level of loans outstanding (equivalent to 10.02 billions in nominal terms) and a decline of 5.05 billions for bonds (equivalent to 8.55 billions in nominal terms), and they are thus remarkably similar.[20] These results suggest that corporate borrowers may take advantage of low interest rates to place new bonds and restructure existing debts, but as interest rates rise and corporate balance sheets improve, companies are expected to explore alternative sources of funding. Thus, C&I loans and bonds are substitutes at the aggregate level. It follows that the development of an efficient corporate bond market is important in at least two respects. First, to reduce the market power of banks. Secondly, to increase the resilience of the economy against bank driven credit crunches. Finally, we find no evidence that changes in interest rates influence the levels of issuance in the equity market.[21]

[19] With regard to C&I loans aggregate, an increase of 0.1% on monthly basis corresponds to an annual increase of 1.21%.

[20] For this calculation we have used the average level of bond issuance and the 3SLS empirical estimates reported in Table 2.

[21] Changes in short-term interest rates have been omitted from the final version of our empirical results as not statistically significant.

V.2 GARCH analysis

To further investigate the relationship among the three different aggregates we make use of the multivariate Diagonal VECH GARCH model of Eqs.(4)-(7) to analyze the statistical properties of their conditional variance/covariance matrix. Such econometric framework makes it possible to investigate the extent of interrelation over time among the three markets. Because shocks in the mean equations are the main actors in GARCH models, it is paramount that the mean equations for the three aggregates are properly specified. Thus, we carry out a preliminary analysis to identify the correct model specifications that will be used in the estimation of our multivariate GARCH. Although we find some evidence of cross interactions among the log first differences of the three aggregates, in order to reduce the number of parameters under estimation we choose to fit the three series with simple AR processes. For each series we identify the optimal lag length by fitting AR(p)-GARCH specifications and by checking for the absence of serial correlation in the standardized and squared standardized residuals.[22] Following this selection criteria we find that the best fitting models are an AR(3)-GARCH(1,1) for STOCKS and an AR(5)-GARCH(1,1) for C&I loans. With regard to the aggregate BONDS the best fitting model is an AR(4) without GARCH effects in the residuals. These specifications are then used to estimate the multivariate Diagonal VECH GARCH model whose empirical estimates are reported in Table 3.[23] The likelihood ratio (LR) test for the null of constant covariance matrix strongly suggests that time varying conditional covariances are important when modelling the covariance matrix $\Sigma(t)$ of the different aggregates.[24] The estimates of the coefficients attached to the product of the shocks $\varepsilon_{i,t-1}\varepsilon_{j,t-1}$ range from 0.043 to 0.269 for the variances, and from 0.194 to -5.04e^{-3} for the covariances. These parameters are all statistically significant at the 5% level except those that govern the conditional covari-

[22] We identify the value of p as the minimum number of lags which ensures standardized and squared standardized residuals not serially correlated.

[23] To save space the empirical estimates of the mean equations (4)-(6) are not reported.

[24] The LR statistic for the null $H_0 : a_{11} = a_{12} = a_{13} = a_{23} = a_{33} = b_{11} = b_{12} = b_{13} = b_{23} = b_{33} = 0$ calculates to 162.9, and with the degrees of freedom being equal to 10, the null is rejected at standard significance levels.

ances. Finally, the estimates for the coefficients attached to lagged variances span from 0.906 for STOCKS to 0.153 for C&I loans, with the latter not statistically significant. The coefficients attached to lagged conditional covariances are all statistically significant at the 5% level except those governing the covariance between BONDS and C&I loans. The null $H_0 : a_{ij} = b_{ij} = 0$ is soundly rejected at standard significance levels for i,j=$\{(2,1),(3,1)\}$ whereas the same hypothesis cannot be rejected for i,j=(3,2).When comparing the conditional volatilities of the different aggregates, past shocks seem somewhat more important for C&I loans. Past volatility, on the other hand, plays a more relevant role for STOCKS. The lower panel of Table 3 reports the diagnostic tests. Ljung-Box Q statistics at lags 4 and 8 for the standardized residuals as well as the LM Breush-Godfrey test at lag 4 suggest negligible degrees of serial correlation. Ljung-Box Q statistics as well as ARCH LM test are then calculated for the standardized squared residuals. Also in this case the two tests provide convincing evidence that our model adequately captures the conditional heteroscedasticity in the second moments.[25]

Figure 1 displays the pairwise conditional correlation coefficients among the three aggregates. They all vary over time and follow significantly different patterns. The conditional correlation coefficient between STOCKS and BONDS is always positive and it fluctuates around a mean value of 0.27. Thus, idiosyncratic shocks in the two markets are positively interrelated. On the contrary, the conditional correlation between C&I loans and STOCKS fluctuates from positive to negative values during the period 1973-2001 whereas for the last six years a clear change in the trend seems to occur as the correlation coefficient progressively takes negative values up to -0.3. Even more interestingly, the conditional correlation between C&I and BONDS is consistently negative over time, with a mean value of -0.15, providing further support to the hypothesis that, for large non financial corporations, the two aggregates are substitutes.

[25]The stationary conditions for the Diagonal Vech model are fulfilled as the largest eigenvalue of the matrix $\mathbf{A}+\mathbf{B}$ is 0.854. This implies that for all the three aggregates the unconditional covariance matrix exists.

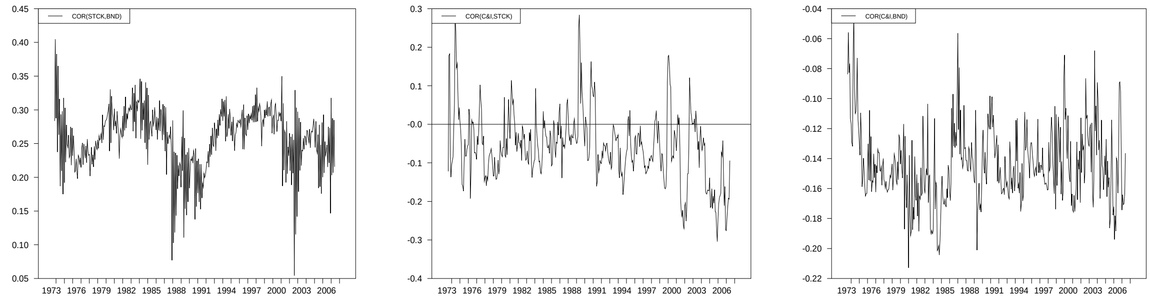

Fig. 1: Conditional correlation coefficients between STOCKS and BONDS (left diagram), C&I loans and STOCKS (centre), and C&I loans and BONDS (right).

VI The Determinants of Linkages Among Markets

The final part of the present analysis involves the identification of potential macroeconomic and financial factors that could explain the time-varying nature of the co-movements among the three markets. We investigate this issue by regressing the time-varying correlations previously estimated against business cycle and monetary policy indicators such as changes in the Industrial Production and Composite Leading Indicators indices, short-term interest rates as well as returns on stock and bond markets. Moreover, we include a set of additional control variables such as aggregate stock market liquidity and development of financial and credit markets. Given that all the above variables have been found to alter returns as well as the levels of co-movement on secondary markets, the same type of indicators might also exert an impact on the time-varying correlations among the volumes of issuance in primary markets.[26] Moreover, in order to analyze the impact of a number of major events on the linkages between the three aggregates considered, we supplement the above regressions with a set of dummy variables that captures episodes such as the 1979 Oil Crisis, the 1981, 1991 and 2001 economic recessions, the stock market crash of October 1987 as well as the burst of the dot-com bubble (10/03/2000-27/09/2002).

As suggested by De Goelj and Marquering (2004), tests for the constancy of the conditional correlation indices can be carried out by estimating AR regressions of the correlation coefficients and by checking the statistical significance of the AR parameters. This

[26]Similar macroeconomic and financial indicators have been used in previous studies on secondary markets. See, e.g., Kim et al. 2005 and Wang and Moore 2008.

Table 3: Maximum Likelihood Estimates of the Diagonal VECH GARCH Model of (4)-(7) for Monthly Volumes of STOCKS, BONDS and C&I loans.

	$\textbf{STOCKS}_t$		$\textbf{BONDS}_t$		$\textbf{C\&I}_t$	
	coeff	**s.e.**	**coeff**	**s.e.**	**coeff**	**s.e.**
$\textbf{const}_1$	0.007^{*}	0.003				
$\textbf{const}_{21}$	0.071^{***}	0.019				
$\textbf{const}_{31}$	$-0.2e^{-4}$	$0.4e^{-4}$				
$\textbf{const}_2$			0.145^{***}	0.010		
$\textbf{const}_{32}$			-0.0003	0.0005		
$\textbf{const}_3$					$1.8e^{-5}$	$8.5e^{-6}$
$\sigma^2_{1,t-1}$	0.906^{***}	0.029				
$\sigma^2_{21,t-1}$	-0.740^{**}	0.295				
$\sigma^2_{31,t-1}$	0.797^{**}	0.294				
$\sigma^2_{2,t-1}$			–	–		
$\sigma^2_{32,t-1}$			0.034	1.696		
$\sigma^2_{3,t-1}$					0.153	0.273
$\varepsilon^2_{1,t-1}$	0.043^{**}	0.016				
$\varepsilon_{2,t-1}\,\varepsilon_{1,t-1}$	-0.016	0.028				
$\varepsilon_{3,t-1}\,\varepsilon_{1,t-1}$	0.048	0.053				
$\varepsilon^2_{2,t-1}$			–	–		
$\varepsilon_{3,t-1}\,\varepsilon_{2,t-1}$			-0.015	0.054		
$\varepsilon^3_{3,t-1}$					0.269^{***}	0.079
$\mathbf{R}^2$		–		–		–
Q(4)	5.200		2.903		2.528	
	(0.157)		(0.406)		(0.470)	
Q(8)	7.420		9.729		5.386	
	(0.386)		(0.204)		(0.612)	
$\textbf{LM(4)}^{\flat}$	8.237		18.04		15.59	
	(0.766)		(0.114)		(0.210)	
$\textbf{Q}^2\textbf{(4)}$	4.286		6.418^{*}		1.799	
	(0.233)		(0.092)		(0.615)	
$\textbf{Q}^2\textbf{(8)}$	8.544		6.658		6.507	
	(0.287)		(0.465)		(0.481)	
$\textbf{ARCH(4)}^{\natural}$	3.687		5.228		1.723	
	(0.450)		(0.264)		(0.786)	

Notes: Sample period 1973:01 - 2007:06. Empirical estimates for parameters of eqs.(4)-(6) not reported. * = significant at 10%, ** = at 5% and *** = at 1%. Adjusted R^2 calculated as $1-(1-R^2)/(T-1/T-k)$. Q(n) and Q^2(n) are Ljung-Box statistics for serial correlation up to lag n in the standardized and squared standardized residuals. $\flat$ LM test for serial correlation in standardized residuals up to lag 4. $\natural$ ARCH LM test for heteroscedasticity in standardized residuals up to lag 4. P-values in parenthesis.

test is of particular importance to investigate whether the conditional correlation between BONDS and C&I loans is actually time varying, as the null $a_{32}=b_{32}=0$ cannot be rejected at standard significance levels. Thus, we supplement our regressions with lagged values of the correlation coefficients.[27] More specifically, we estimate the following linear

[27] The inclusion of lagged values of the conditional correlation coefficients is consistent with previous studies such as Kim et al. (2005) and Wang and Moore (2008).

regressions (for i,j=1,2,3 and i$\neq$j):

$$\rho_{ij,t} = a_0 + \sum_{l=1}^{L} a_l \rho_{ij,t-l} + b_1 \Delta CLI_t + b_2 \Delta IPI_t + b_3 \Delta i_t + b_4 R_{S\&P,t} + \tag{9}$$

$$+ b_5 R_{BCBI,t} + b_6 FD_{1,t} + b_7 FD_{2,t} + b_8 V_t + \sum_{i=1}^{6} c_i DUM_{i,t} + \varepsilon_{ij,t}$$

$$\sigma^2_{ij,t} = \omega + \alpha \sigma^2_{ij,t-1} + \beta \varepsilon^2_{ij,t-1} \tag{10}$$

where $\rho_{ij,t-l}$ is the lagged dependent variable, $FD_{1,t}$ and $FD_{2,t}$ are, respectively, the ratios of stock market capitalization and total loans to GDP which are commonly used to capture the stock market and banking sector development, and V_t is the logarithm of the stock market turnover by volume to proxy the liquidity of financial markets.[28] The dummy variables $DUM_{i,t}$ capture the 1979 Oil Crisis (i=1), the 1981, 1991, 2001 economic recessions (i=2,4 and 5), the Stock Market Crash of October 1987 (i=3) and the dot-com bubble (i=6).[29] All data are monthly observations that span from 1973:06 to 2007:06.

Maximum Likelihood estimates of eqs.(9)-(10) are reported in Table 4. Diagnostic tests indicate that the explanatory power of the regressions is fairly high as the R-squared span from 0.14 to 0.84.[30] Moreover, Ljung-Box Q-statistics applied to standardized and squared standardized residuals clearly indicate absence of serial correlation and conditional heteroscedasticity. Tests for the constancy of the conditional correlations clearly suggest that all the estimated coefficients are not constant over time and characterized by persistence, as all the AR parameters are statistically significant at standard levels.

The conditional correlation between STOCKS and C&I fluctuates from positive to negative for large parts of the sample, suggesting that the two markets are loosely interrelated.

[28] The number of lags L for the regressions with dependent variables $\rho_{12.t}$ $\rho_{13,t}$ and $\rho_{23,t}$ are, respectively, 2, 2 and 3.

[29] The dummy variables assume value one when the episode occurs, keep such value for the next 24 observations, and then exponentially decay toward zero with a factor of $0.95^{(t-t^*)}$ (where t^* is the date a specific episode takes place). With a decaying factor equal to 0.95 the half-life is 14 months. Dummy variables so constructed make it possible to avoid problems of collinearity when simultaneously included in the same regression.

[30] The large explanatory power is not the result of spurious regression since the Augmented Dickey-Fuller tests in Table 4 indicate that the standardized residuals are stationary.

This moderate co-movement is exclusively driven by expectations on the cycle and monetary policy. More specifically, the weak negative linkage between these two markets strengthens whenever increases in the Composite Index of Leading Indicators (signalling expectations for a better economic outlook) and short-term interest rates (signalling potential overheating of the economy) occur.

The correlation coefficients between the primary markets for shares and bonds, and between C&I loans and bonds seem the most interesting relationships, as they clearly assume values different from zero over time, and they can be explained by the macroeconomic and financial indicators previously defined. The broad brush picture we obtain is that the levels of co-movement between the issuance of bonds and C&I loans mainly respond to stock and bond markets valuations, whereas both the above macro and financial factors appear to play an important role in explaining the co-movement between primary markets of shares and debt.

Current business conditions, as captured by fluctuations in the Industrial Production Index, exert a negative impact on the linkage between STOCKS and BONDS. Thus, in periods of economic growth the two markets become less interrelated. This dynamic can be explained by the result that the issuance of bonds is counter cyclical, whereas the issuance of shares does not respond to the industrial production.[31] Expectations of future economic growth, as captured by higher levels of the Composite Index of Leading Indicators, foster the positive linkage between STOCKS and BONDS, whereas they do not exert any impact on the co-movements between bond and credit markets. For instance, a joint increase in both the issuance of shares and bonds might occur when, in the middle of an economic slowdown where the issuance of bonds increases, expectations for future economic activity arise so that financial markets respond with an increase in the volume of shares issued.

Bond market valuations is an important driving force for the correlation between BONDS and C&I loans. Higher levels of the BCBI (implying lower yields) foster the

[31] See the empirical estimates reported in Table 2.

negative correlation between the two aggregates. This result suggests that, whenever bonds valuation improves, industrial firms find it more economical to place new bonds on the primary market. At the aggregate level, this would eventually increase the overall volume of bonds issued, and reduce the demand for C&I loans. Both stocks and bonds valuations exert strong influence also on the linkage between the volumes issued of shares and debt. Higher share prices make the positive correlation between the two primary markets stronger, whereas increases in the BCBI index weaken such linkage. In this case, the net effect generated by the two indices on the correlation coefficient is, on average, positive and equal to 0.08. We can make sense of this result by recalling that stock prices measure future expected dividends properly discounted, whereas the bond index proxies the discount factor. Thus, lower levels in the BCBI index, indicating higher interest rates, must be associated with higher discounting of future profits and lower levels of the stock index.[32] Moreover, higher expected dividends are associated with higher future levels of economic activity and interest rates. Thus, the opposite sign of the coefficients attached to the two indices suggests that expectations of higher future dividends increase the correlation, whereas lower interest rates reduce it. As higher interest rates are associated with higher expected inflation, we can conclude that the above correlation increases whenever real or nominal expected dividends rise.

Both higher ratios of credit and stock market capitalization to GDP weaken the negative correlation between BONDS and C&I loans as well as the positive correlation between STOCKS and BONDS, suggesting that the above linkages become blurred in periods of credit expansion and market booms, and stronger in periods of credit contraction and bearish stock markets. On the one hand, these results are against the hypothesis of financial contagion between bond and credit markets, as we find that the substitution between the two aggregates strengthens significantly when the above ratios decrease, so that the two sources of finance do not dry up simultaneously in periods of financial distress.

[32]Returns on share and bond indices are, in fact, positively correlated for the sample period under analysis. More specifically, a 1% increase in the return of the S&P500 implies an increase in the return of the BCBI of the order of 0.15%.

On the other hand, the above results support the hypothesis of contagion between primary markets of shares and debt as the co-movement between the two aggregates becomes stronger in periods of reduced availability of credit.

Similarly, the level of liquidity in the secondary stock market exerts a statistically significant impact with expected positive sign on the correlation between the volumes of BONDS and C&I loans, whereas the impact is negative for the correlation between the volumes of shares and bonds issued. In both cases our empirical estimates suggest that, as the stock market becomes more liquid, the costs attached to the issuance of new shares decrease, and more options become available to firms in search of capital. This, in turn, would weaken the relationship of substitution between BONDS and C&I loans, and that of complementarity between the primary markets of shares and bonds. Finally, fluctuations in the short-term interest rate do not exert any significant impact on the levels of both the above correlation coefficients.

It is also evident that the macroeconomic and financial shocks, as captured by the dummy variables for the 1981, 1991 economic recessions, the 1987 stock market crash and dot-com bubble, exert statistically significant impacts by weakening the linkages among the three markets. This result, in turn, provides convincing evidence against the hypothesis of financial contagion. For instance, the occurrence of the 1987 stock market crash weakens the co-movement between primary markets of shares and bonds, and between those of shares and C&I loans . Other things being equal, the occurrence of the crash reduces the correlation between STOCKS and BONDS by more than 10%.[33] Moreover, both the 1991 recessions and the burst of the dot-com bubble weaken the co-movement between the primary markets of shares and debt, whereas the 1981 recession exerts a weakening impact on the levels of co-movement between the primary markets of stocks and bonds, and C&I loans. For instance, the occurrence of the 1981 recession causes a reduction in the levels of co-movement between BONDS and C&I loans of the order of 20%.

[33] Similar reductions in the levels of co-movement occur for the correlation between the volumes of STOCKS and C&I loans.

Table 4: Maximum Likelihood Estimates of Eqs.(9)-(10).

	ρ_t(C&I,STOCKS)		ρ_t(C&I,BONDS)		ρ_t(STOCKS,BONDS)	
	coeff	s.e.	coeff	s.e.	coeff	s.e.
ρ_{t-1}(C&I,STOCKS)	0.902***	0.048	–	–	–	–
ρ_{t-2}(C&I,STOCKS)	-0.130***	0.048	–	–	–	–
ρ_{t-1}(C&I,BONDS)	–	–	0.146***	0.052	–	–
ρ_{t-2}(C&I,BONDS)	–	–	0.074	0.049	–	–
ρ_{t-1}(STOCKS,BONDS)	–	–	–	–	-0.151***	0.044
ρ_{t-2}(STOCKS,BONDS)	–	–	–	–	0.627***	0.026
ρ_{t-3}(STOCKS,BONDS)	–	–	–	–	0.159***	0.021
$\Delta\mathbf{CLI}_{1,t}$	-1.572***	0.552	0.170	0.299	0.952***	0.294
$\Delta\mathbf{IPI}_{1,t}$	0.371	0.361	-0.150	0.179	-0.377***	0.163
$\mathbf{i}_{1,t}$	-0.009*	0.005	-0.002	0.002	0.001	0.002
$\mathbf{R}_{S\&P,t}$	-0.041	0.058	0.050	0.031	0.101***	0.026
$\mathbf{R}_{BCBI,t}$	-0.042	0.123	-0.137**	0.067	-0.132**	0.056
$\mathbf{FD}_{1,t}$	-0.0003	0.004	0.005**	0.002	-0.0005	0.002
$\mathbf{FD}_{2,t}$	0.0012	0.0008	0.001**	0.0004	-0.003***	0.0005
$\mathbf{V}_t$	-0.0002	0.005	0.004*	0.002	-0.018***	0.003
$\mathbf{DUM}_{1,t}$	-0.021	0.014	-0.0002	0.007	0.008	0.006
$\mathbf{DUM}_{2,t}$	0.032**	0.016	0.031***	0.009	0.0004	0.007
$\mathbf{DUM}_{3,t}$	0.026**	0.011	0.001	0.006	-0.036	0.007
$\mathbf{DUM}_{4,t}$	0.011	0.015	0.002	0.008	-0.039***	0.010
$\mathbf{DUM}_{5,t}$	0.012	0.015	0.005	0.008	0.018	0.011
$\mathbf{DUM}_{6,t}$	-0.023	0.019	-0.008	0.009	-0.028**	0.013
ω	–	–	0.0003***	0.0001	0.0004***	0.0001
α	–	–	0.209***	0.075	0.542***	0.157
β	–	–	0.299*	0.171	–	–
ADF	-7.019	(0.000)	-6.987	(0.000)	-5.999	(0.000)
Q(4)	2.184	(0.535)	2.564	(0.463)	5.101	(0.164)
Q(8)	4.584	(0.711)	4.695	(0.697)	6.460	(0.487)
$\mathbf{Q}^2$**(4)**	6.915	(0.075)	4.661	(0.198)	6.162	(0.104)
$\mathbf{Q}^2$**(8)**	10.02	(0.187)	6.589	(0.473)	7.402	(0.388)
$\mathbf{R}^2$	0.759	-	0.142	-	0.842	-

Note: Sample periods 1973:09 - 2007:06. Estimates of constant term for the three regressions not reported. * = significant at 10%, ** = at 5% and *** = at 1% level. Standard errors in parenthesis. ADF is the Augmented Dickey-Fuller test applied to standardized residuals with lags 8. Q(n) and Q^2(n) are Ljung-Box statistics for serial correlation in standardized and squared standardized residuals up to lag n. P-values in parenthesis. Adjusted R^2 calculated as $1-(1-R^2)/(T-1/T-k)$.

VII Conclusion

In this paper we investigate the short-run interrelations among the volumes of securities placed in the primary markets of shares and corporate bonds, and those of C&I loans by using simultaneous equations and multivariate GARCH approaches. The broad brush picture we obtain suggests that the equilibrium aggregate volumes of the different securities are interrelated, and that macroeconomic indicators which proxy business conditions explain sizeable proportions of their variability. The estimated conditional correlations provide convincing evidence of the changing levels of co-movement in the three aggregates over time.

We find that the issuance of stocks and bonds are mutually reinforcing processes. This result supports the argument put forward by Lowry (2003) that the issuance of these two types of securities can be driven by common factors. Furthermore, in line with the literature, we find that the issuance of both shares and bonds is forward-looking, as they strongly react to current stock and bond market valuations. This can reflect either a valuation of the expected profitability of investments, or an opportunistic timing of the issuance on part of insiders, in order to exploit asymmetric information.[34]

Moreover, our results suggest that primary placements of corporate bonds and C&I loans are interrelated, with the causal relationship being unidirectional as C&I loans granger-cause the issuance of bonds. We find that the two types of securities respond in opposite ways to changes in current levels of economic activity, as well as monetary policy stance. Higher levels of industrial production are, in fact, associated with larger issuance of C&I loans, probably because industrial firms finance inventories by resorting to unused commitment loans, and lower levels of bond issuance. This finding is supported by the result that the conditional correlation between the two aggregates assumes negative values over the entire sample under analysis. A plausible explanation is that in periods of expansion the cyclical component of the supply declines, as firms benefit from larger cash flows and need less external finance. Moreover, bond volumes quickly increase with the implementation of expansionary monetary policy whereas bank C&I lending rises when the monetary policy is tightened. Thus, as a recession lurks, industrial firms need to rely on bank lending because financial markets freeze, but as the FED reacts by slashing interest rates, corporate bonds become cheaper, even if risk premia remain substantial, so that firms proportionally reduce their recourse to bank lending. As a consequence, bond

[34] The issuance of shares is also positively affected by past stock market valuations, indicating that market momentum, rather than just the level of the price index, is a strong factor behind primary placements.

markets may be very important to mitigate the impact of a bank-driven credit crunch.

The empirical estimation of the conditional correlation coefficients reveals that the co-movements among the three markets can be explained by business cycle and monetary policy indicators such as short-term interest rates, Industrial Production and Composite Index of Leading Indicators. The levels of credit and stock market capitalization to GDP have mixed impacts on the levels of interrelation among the three markets. On the one hand, our results suggest that the primary markets for corporate bonds and C&I loans do not dry up simultaneously in periods of financial distress. On the other, lower levels of credit to GDP strengthen the linkage between primary markets of shares and debt. The former result provides evidence against the hypothesis of financial contagion, whereas the latter seems to support this hypothesis. Finally, macroeconomic and financial shocks such as the 1981 and 1991 economic recessions as well as the 1987 stock market crash and the burst of the dot-com bubble consistently weaken the levels of interrelation among the three markets, providing substantial evidence against the hypothesis of contagion. This last result is particularly neat for the primary markets of shares and bonds, and it suggests that the impact of the above exogenous shocks on real variables is mitigated by the evidence that industrial firms can obtain finance from different sources that become less interdependent in periods of financial distress or economic downturn.

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References

Chami, Ralph and Thomas Cosimano. “Monetary Policy with a Touch of Basel.” *Journal of Economics and Business* 62 (2010), 161–175.

Connolly, R. A., C. Strivers, and Sun L.. “Commonality in the Time-Variation of Stock-

Stock and Stock-Bond Return Comovements." *Journal of Financial Markets* 10 (2007), 192–218.

Corsetti, Giancarlo, Marcello Pericoli, and Sbracia Massimo. "Some Contagion, Some Interdependence: More Pitfalls in Tests of Financial Contagion." *Journal of International Money and Finance* 24 (2005), 1177–1199.

Cosimano, Thomas. "The Banking Industry under Uncertain Monetary Policy." *Journal of Banking & Finance* 12 (1988), 117–139.

Den Haan, Wouter J., Steven W. Sumner, and Guy M. Yamashiro. "Bank loan portfolios and the monetary transmission mechanism." *Journal of Monetary Economics* 54 (2007), 904–924.

Dittmar, Amy and Anjan Thakor. "Why Do Firms Issue Equity?" *The Journal of Finance* 62 (February 2007), 1–54.

Dorn, Daniel. "Does Sentiment Drive the Retail Demand for IPOs?" *Journal of Financial and Quantitative Analysis* 44 (2009), 85–108.

Gertler, Mark and Simon Gilchrist. "The cyclical behaviour of short-term lending." *European Economic Review* 37 (1993), 623–631.

Goelj, Peter De and Wessel Marquering. "Modelling the Conditional Covariance Between Stock and Bond Returns: A Multivaruiate Approach." *Journal of Financial Econometrics* 2 (2004), 531–564.

Granger, Clive and Paul Newbold. "Spurious Regressions in Econometrics." *Journal of Econometrics* 2 (1974), 111–120.

Helwege, Jean and Nellie Liang. "Initial Public Offering in Hot and Cold Markets." *Journal of Financial and Quantitative Analysis* 39 (2004), 541–560.

Hovakimian, Armen, Tim Opler, and Sheridan Titman. "The Debt-Equity Choice." *Journal of Financial and Quantitative Analysis* 36 (March 2001), 1–24.

Ivanov, Vladimir and Craig M. Lewis. “The Determinants of Market-Wide Issue Cycles for Initial Public Offerings.” *Journal of Corporate Finance* 14 (2008), 567–583.

Kashyap, Anil, Jeremy C. Stein, and David W. Wilcox. “Monetary Policy and Credit Conditions: Evidence for the Composition of External Finance.” *The American Economic Review* 83 (1993), 78–98.

Kim, Suk J., Fariboz Moshirian, and Eliza Wu. “Dynamic Stock Market Integration Driven by the European Monetary Union: An Empirical Analysis.” *Journal of Banking and Finance* 29 (2005), 2475–2502.

Lowry, Michelle. “Why Does IPO Volume Fluctuate So Much?” *Journal of Financial Economics* 67 (2003), 3–40.

Mayfield, Scott E. “Estimating the Market Risk Premium.” *Journal of Financial Economics* 73 (2004), 465–496.

Pagano, Marco, Fabio Panetta, and Luigi Zingales. “Why Do Companies Go Public? An Empirical Analysis.” *The Journal of Finance* 53 (February 1998), 27–64.

Pástor, Ĺubos and Pietro Veronesi. “Rational IPO Waves.” *The Journal of Finance* 60 (2005), 1713–1757.

Pericoli, Marcello and Massimo Sbracia. “A Primer on Financial Contagion.” *Journal of Economic Surveys* 17 (2003), 571–608.

Ritter, Jay R. and Ivo Welch. “A Review of IPO Activity, Pricing and Allocations.” *The Journal of Finance* LVII (August 2002), 1795–1828.

Schwert, William G.. “Tests for Unit Roots: A Monte Carlo Investigation.” *Journal of Business and Economic Statistics* 7 (1989), 147–158.

Syllignakis, Manolis N. and Georgios P. Kouretas. “Dynamic Correlation Analysis of Financial Contagion: Evidence from the Central and Eastern European Markets.” *International Review of Economics and Finance* 20 (2011), 717732.

Voronkova, S. “Equity Market Integration in Central European Emerging Markets: a Cointegration Analysis with Shifting Regimes.” *International Review of Financial Analysis* 13 (2004), 633647.

Wang, Ping and Tomoe Moore. “Stock Market Integration for the Transition Economies: Time-Varying Conditional Correlation Approach.” *Manchester School* (2008), 116–133.