“Pure contagion vs. financial interconnection in the subprime crisis context: Short- and long-term dynamics”

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PURE CONTAGION VS. FINANCIAL INTERCONNECTION IN THE SUBPRIME CRISIS CONTEXT: SHORT- AND LONG-TERM DYNAMICS

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Abstract

This paper examines the difference between pure contagion and financial interconnection by studying the U.S. and some American and Asian markets in the subprime crisis context. These markets are affected by the mortgage crisis, with data available from January 1, 2003 to December 30, 2011. The paper first identifies the turmoil period via the wavelet technique and adopts cointegration and Granger causality approaches by estimating vector autoregressive (VAR) and vector error correction models (VECM) models. Based on daily returns from stock market indices in five American countries (Mexico, Brazil, Canada, Argentina, and the U.S.) and eight Asian ones (Hong Kong, Japan, India, Indonesia, Malaysia, Singapore, Korea, and China), the results show eight cases of pure contagion and 10 cases of financial interconnection. In addition, there were high co-movements in the short term and low co-movements in the long term for financial interconnection cases. These findings have several implications for investors looking to diversify their portfolios internationally and for portfolio managers to expect and limit market risk. The results provide additional guidance to regulators and policymakers.

Keywords subprime crisis, pure contagion, financial interconnection, cointegration, Granger causality

JEL Classification G15, C58

INTRODUCTION

During the last few years, crises have followed one another and multiplied, such as the Asian crisis (1997), the Technological crisis (2002), the Subprime crisis (2007–2009), the Sovereign debt crisis, and COVID-19. Given the links that exist between different countries, a crisis affecting a given country can then spread through the international financial markets and affect others. International finance uses the term "contagion" to describe this type of effect.

The links between markets, especially the phenomenon of financial contagion during crises, have received growing attention from academicians. Indeed, since the 1997 Asian crisis, contagion has become the most discussed topic in financial markets (Eichengreen et al., 1996; Forbes & Rigobon, 2001). These studies investigate the existence of contagion, taking into account various political, economic, and financial factors.

Despite the multitude of research on financial contagion, there is no consensual definition and measure of contagion, nor for the channels of crisis transmission. Understanding the notion of financial contagion is important to identify the mechanisms that have led to such

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propagation and to identify the existing relationships between markets. Particularly, interconnection and pure contagion are important parts of financial stability and risk assessment of the financial system of a country. According to Forbes and Rigobon (2002), fundamental contagion (or interconnection) is based on the effects of interconnections through trade, economic, and financial links between countries. However, pure contagion is related to rational or irrational investors' behavior, such as herd behavior, financial panics, increased risk aversion, and loss of confidence. The "pure contagion" phenomenon means that the financial crisis is linked to the actions of investors, not to changes in fundamentals or macroeconomic indicators as with fundamental contagion.

The extant literature has well investigated financial contagion and the associated volatility. The paper extends prior literature in the context of the subprime crisis to detect cases of both interconnection and pure contagion using cointegration and Granger causality approaches. It first identifies the turmoil period via the wavelet technique and then adopts cointegration and Granger causality approaches. This method provides information on whether pure contagion and financial interconnection exist in the short and long terms. In addition, it indicates the causality direction between returns from different markets.

1. LITERATURE REVIEW

The mere definition of contagion remains controversial. Eichengreen et al. (1996) define contagion as "a significant increase in the likelihood of a crisis in one country, conditional on the occurrence of a crisis in another country". Forbes and Rigobon (2001) argue that financial contagion is "a significant increase in cross-market linkages after a shock to an individual country (or group of countries)." In fact, financial contagion exists when the degree of co-movements between two markets is high during the stability period and keeps increasing after the crisis.

Empirically, a large body of literature has investigated financial contagion and how to measure stock market contagion and the associated volatility. Baur (2012) examines the contagion among 25 major stock markets and their real economy sectors during the global financial crisis. He proves that no single country or sector was immune to the adverse effects of the crisis, though some sectors were affected less severely. Samarakoon (2011) investigates the transmission of shocks between the U.S. and foreign markets. The author shows an interdependence and a contagion in emerging markets, with important regional variations. Kenourgios and Dimitriou (2015) examine 10 sectors in six developed and emerging regions' markets during different phases of the crisis, testing different channels of financial contagion using dynamic conditional correlation from the multivariate Fractionally Integrated Asymmetric Power ARCH model. The authors find contagion effects in the subprime crisis across regional stock markets and financial and non-financial sectors. Recently, Zorgati and Lakhal (2020) examined the spatial dimension's influence on financial contagion in the context of the subprime crisis. Using a local correlation measure, the authors prove the existence of spatial contagion between the U.S. market and others in the American region. As for markets that are geographically distant from the U.S., the authors have proved that spatial contagion exists between only some groups of countries. In addition, Davidson (2020) analyzes contagion and interconnection in Latin America using a novel model-switching approach and shows the importance of macroeconomic and uncertainty channels.

In a different context of the Eurozone Sovereign Debt crisis, Kalbaska and Gątkowski (2012) analyze the dynamics of the credit default swap market for some European countries. Using the Granger-causality test and the Exponentially Weighted Moving Average (EWMA) correlation analysis, they find the existence of a contagion effect. In the same context, Kenourgios (2014) investigates volatility contagion across U.S. and European stock markets during the Global Financial Crisis and the Eurozone Sovereign Debt Crisis. The author uses asymmetric conditional correlation dynamics across stable and crisis periods, as well as across different phases of both crises, to support the existence of contagion in cross-market volatilities.

Gharib et al. (2020) investigate the bilateral contagion effects across oil and gold using time-varying Granger causality tests. They find contagion effects from oil and gold market bubbles during the 2014–2015 crash and, lately, the COVID-19 pandemic. In the same vein, Akhtaruzzaman et al. (2021) examine the occurrence of financial contagion through financial and non-financial firms during the COVID-19 period. They find that China and Japan transmitted more spillovers than they received during the COVID-19 period and show that the hedging costs increased during the COVID-19 period to optimize portfolios.

To detect the long-term relationships between time series, several studies have used the cointegration theory. Tan (1998) emphasizes the importance of the contagion phenomenon on the stock market during the Asian crisis. The author studies stock market indices' co-movements based on the error correction model (ECM) and shows that a significant change occurs in stock market indices' co-movements during stable and crisis periods. Masih and Masih (1999) use the cointegration technique on four Asian countries' stock indices and also detected contagion. Yang et al. (2005) study short- and long-term cointegration relationships between the United States, Japan, and 10 other Asian markets. Their results indicate that strong integration exists during the Asian crisis and that this integration was exacerbated after the crisis. Dewandaru et al. (2016) investigate contagion among Asia-Pacific equity markets (Japan, Australia, and Hong Kong) during 12 major crises using discrete and continuous wavelet decompositions. They document a contagion effect of subprime crisis to fundamental links, and that the Japanese market played a dominant role. Gómez-Puig and Sosvilla-Rivero (2016) assess the transmission of the European sovereign debt crisis, applying a dynamic Granger-causality approach to detect contagion. Using a logit model to distinguish between pure and fundamental contagion, the authors find the coexistence of pure and fundamentals-based contagion.

More recently, Ozparlak (2020) investigates the long- and short-run impact of the COVID-19 crisis on CDS markets and stock markets using the Cointegration methodology. He finds a long-term relationship between the total cases of COVID-19 and China, France, the United Kingdom, Germany, Turkey, and Spain. However, there is no long-term relationship between the total cases of COVID-19 and Italy and the USA.

This article attempts to examine the difference between pure contagion and financial interconnection between the U.S. market and other American and Asian markets in the context of the subprime crisis.

2. METHODOLOGY

The paper uses the daily series of stock indices from American and Asian countries (five days a week).[1] Indeed, Bannigidadmath and Narayan (2016) argue that daily data are better than monthly ones as they include richer information. Considering daily data is also helpful, given the sufficient number of observations to test our hypotheses.

The U.S. subprime mortgage crisis of 2007–2010 affected all these countries. The studied period covers nine years, from January 1, 2003 to December 30, 2011. The sample is divided into two groups of countries. The first is the Asian region, namely Hong Kong (Hang Seng HSI), Japan (Nikkei 225), India (BSESN), Indonesia (JKSE), Malaysia (KLSE), China (China Shanghai Composite Index (SSE), Korea (KS11), and Singapore (STI). The second group includes American countries, namely Brazil (BVSP), Mexico (MXX), Argentina (Merv), Canada (S&P/TSX), and the U.S. (S&P 500).[2]

Table 1 shows descriptive statistics of stock return indices for different markets during the whole period. The average stock returns indexes are positive and close to zero for all markets except for Japan. These stock returns vary between –0.0006 (Japan) and 0.093 (Indonesia). In addition, the skewness of returns is close to zero and negative for the majority of stock index returns, indicating a low asymmetry in returns. Table 1 also shows that the kurtosis value is over 3, indicating the

1 We choose the countries belonging to the region of America and Asia which data were available from January 1, 2003, to December 30, 2011.When data cannot be obtained due to holidays, bank holidays, or other reasons, the price of the stock index is viewed as equal to the price of the previous trading day.

2 We should note that the data were obtained from this website: http://fr.finance.yahoo.com/

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Table 1. Statistics summary during the total period

Markets	Japan	Hong Kong	India	Indonesia	Malaysia	Korea	China
n.obs	2345	2345	2345	2345	2345	2345	2345
min	–14.343	–13.582	–11.809	–10.953	–12.966	–11.172	–34.959
max	13.234	13.406	15.989	7.759	12.791	13.209	35.03
mean	–0.0006	0.029	0.064	0.093	0.036	0.045	0.020
stdev	1.598	1.6814	1.677	1.497	0.928	1.565	2.042
Skewness	–0.565	0.1492	–0.009	–0.621	–0.719	–0.299	–0.188
Kurtosis	12.352	11.722	7.962	6.839	42.920	8.383	80.096
J.B	15064.6	13464.7	6209.8	4733.50	18052.3	6918.64	627966.6
Q(10)	53.2141***	107.365***	96.2371***	62.9446***	26.331	61.3317***	129.86***
Markets	**Singapore**	**Brazil**	**Argentina**	**Mexico**	**Canada**	**U.S.**	
n.obs	2345	2345	2345	2345	2345	2345	
min	–18.685	–19.979	–35.825	–7.266	–11.731	–9.469	
max	9.734	15.4728	13.953	11.111	10.973	10.42	
mean	0.037	0.077	0.082	0.078	0.0301	0.0157	
stdev	1.353	1.9005	2.068	1.381	1.262	1.335	
Skewness	–1.462	–0.438	–2.410	0.235	–0.725	–0.166	
Kurtosis	24.346	10.656	42.664	6.398	14.576	9.789	
J. B	58866.07	11194.22	180457.2	4031.844	21008.44	9396.3	
Q(10)	81.9579***	66.7625***	30.3927	44.0829**	170.064***	220.638***	

Notes: The Jarque-Bera test is used to check whether the return distribution is normal. The Box-Pierce-Ljung statistic, Q (10) statistic is distributed as a χ2 with 10 degrees of freedom. *, **, and *** are significance levels at 10%, 5%, and 1%, respectively.

non-normality of the return series and the occurrence of extreme values. Jarque-Bera's statistic shows that stock market index returns do not follow a normal distribution, whereas the Box Pierce Ljung portmanteau test of order 10 shows that most index returns are uncorrelated.

2.1. Description methods

The paper uses the cointegration and Granger causality approach methodology, following Sander and Kleimeier (2003) who identify contagion as an increasing number of cointegration relationships between stable and crisis periods. The paper also uses Granger's (1969) causality test to examine relationships between countries.

The cointegration approach, presented by Granger (1983) and Engle and Granger (1987) is considered one of the most importantccon-cepts inceconometrics and time series analyses. It helps detecting the long-term relationship between two or more timecseries. The method selection for data analysis is based on the unit root test results for the stationarity of the variables. Methods commonly used to analyze the stationary time series cannot be used to analyze non-stationary series.

The Engle-Granger and Johansen cointegration technique requires that all series are integrated of the same order 1. Furthermore, the Johansen cointegration methodology needs a large sample size for validity which is not required under other cointegration approaches such as the autoregressive distributed lag (ARDL) developed by Pesaran et al. (2001). The ARDL can be applied regardless of the stationary properties of the variables and has robust results for the cointegration analysis of small and finite sample sizes.

The procedure can be summarized in four steps:

Step 1. Stationarity of series
Step 2. Cointegration test
Step 3. Estimation of the VAR and VECM models
Step 4. Causality of Granger test

3. RESULTS AND DISCUSSION

3.1. Turmoil period identification

This study first identifies the starting date of the subprime crisis using the wavelet technique on the series of stock market returns' indices in the U.S. market – the initiating crisis market.

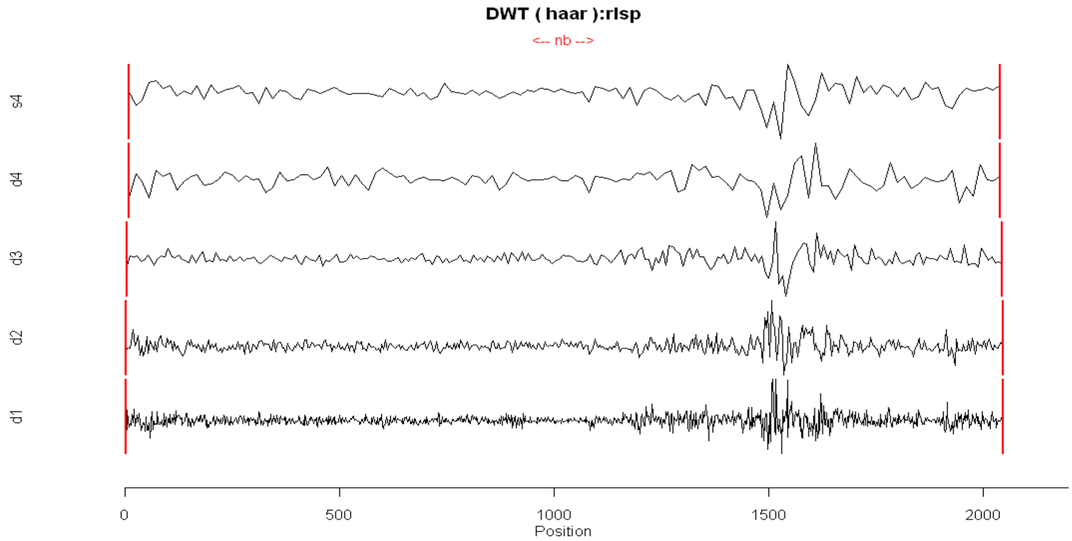

Figure 1. Decomposition of the Haar wavelet's Order 4 (U.S. index (S&P 500))

Figure 1 shows the dissociation of the U.S. index (S&P 500) according to the Haar wavelet at Level 4, indicating that the first sign of crisis appeared on July 17, 2007. The result is consistent with the U.S. Central Bank (Fed), which has stated that the subprime crisis started in the summer of 2007.

To maintain the same crisis period for all the studied markets, we assumed that the mortgage crisis began on August 1, 2007. For more robustness of the results, the study also uses the date of July 17, 2007 identified using the wavelet technique. The results remain unchanged throughout the study.

The paper examines the changes and causal relationships between the different stock markets in the sample using the cointegration technique. It starts by examining the stationarity of the series of returns from the stock market indices for both sub-periods (stability and crisis periods).

3.2. Unit root test

The graphs on the series of returns from the stock market indices presented in Appendix A support the stationarity hypothesis. To confirm these intuitions, stationarity is tested by applying the augmented Dickey-Fuller (ADF) test (Dickey & Fuller, 1981). For more robustness, a more recent stationary test is implemented, which is the General Least Squared Dickey-Fuller (DF-GLS) unit root test. Elliott et al. (1996) showed that DF-GLS unit root test has a greater power than the previous versions of the augmented Dickey-Fuller test. Table 2 shows the results from the unit root ADF tests and DF-GLS test, indicating that all series of stock-level indices were non-stationary during the stability and crisis periods. Indeed, the ADF and DF-GLS tests' values are higher than the critical value of 5%. It is also noticed that all the series

Table 2. ADF and DF-GLS tests of stock index series: stability period (crisis period)

Countries	ADF test		DF-GLS test		I(d)
	Level	First difference	Level	First difference	
JAPAN	–2.6004	–24.804*	–2.3402	–15.6522*	I(1)
	(–1.420)	(–25.725*)	(0.4447)	(–15.890*)	(I(1))
HONG K	–1.619	–24.446*	–1.6752	–10.947*	I(1)
	(–0.606)	(–24.358*)	(–1.496)	(–13.790*)	(I(1))
INDIA	–2.238	–25.628*	–1.3682	–11.546*	I(1)
	(–0.274)	(–23.025*)	(–1.505)	(–10.127*)	(I(1))
INDONESIA	–0.392	–24.166*	–0.4427	–4.1256*	I(1)
	(0.847)	(–21.187*)	(0.1495)	(–15.898*)	(I(1))

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Table 2 (cont.). ADF and DF-GLS tests of stock index series: stability period (crisis period)

Countries	ADF test		DF-GLS test		I(d)
	Level	First difference	Level	First difference	
MALAYSIA	–0.708	–22.1738*	–1.0511	–2.14*	I(1)
	(0.356)	(–25.049*)	(–0.834)	(–13.787*)	(I(1))
KOREA	–1.354	–24.8185*	–0.9389	–7.0274*	I(1)
	(–0.317)	(–23.842*)	(–1.356)	(–7.6931*)	(I(1))
CHINA	2.143	–25.265*	0.847	–3.1012*	I(1)
	(–1.433)	(–27.720*)	(–0.023)	(–3.541*)	(I(1))
SINGAPORE	–1.5528	–24.917*	–1.3928	–9.6267*	I(1)
	(–0.411)	(–23.318*)	(–0.895)	(–14.973*)	(I(1))
BRAZIL	–2.3127	–25.094*	–1.8938	–3.5899*	I(1)
	(0.1736)	(–24.367*)	(–1.348)	(–3.9099*)	(I(1))
ARGENTINA	–3.3224	–24.314*	–2.336	–9.2809*	I(1)
	(–1.944)	(–22.615*)	(–0.478)	(–7.6132*)	(I(1))
MEXICO	–1.864	–24.564*	–0.9123	–4.9128*	I(1)
	(0.427)	(–23.928*)	(–0.910)	(–4.3954*)	(I(1))
CANADA	–3.2548	–23.934*	–2.0068	–2.6423*	I(1)
	(–0.312)	(–26.432*)	(–1.198)	(–17.016*)	(I(1))
U.S.	–3.1215	–26.172*	–2.6949	–2.3832*	I(1)
	(–0.577)	(–26.522*)	(–0.697)	(–3.7428*)	(I(1))

Notes: * is statistical significance at the 5% level.

are integrated at order 1 at a confidence level of 5% and that all the series in the first difference are stationary.

For the period of stability and crisis in level and in the first difference, the critical value is –1.95 at the level 5%.

3.3. Cointegration test

All the series in the study are integrated from order 1[3]. The studied period covers nine years, from January 1, 2003 to December 30, 2011, encompassing the pre-crisis (1,193 observations), and crisis periods (1,150 observations). Therefore, the paper runs a bivariate cointegration test following Engle and Granger (1987). Then, to ensure the robustness of the results from the Engle-Granger cointegration test, we also used Johansen's (1988) cointegration test.

3.3.1. *Bivariate cointegration test using the Engle-Granger methodology (1987)*

Table 3 reports the results from Engle and Granger bivariate cointegration tests for the two sub-periods: crisis and stability. It presents the results from a pairwise cointegration relationship between the U.S. (a country originating from the subprime crisis) and the other countries in our study. Table 2 shows that the number of cointegration relationships between the U.S. and the rest of the sampled countries is greater during the crisis than during the stability period. Indeed, there are 20 cointegration relations in a stable period compared with 17 during a crisis period.

The study concludes with the existence of cointegration relationships between the U.S. and the rest of the sampled countries.

3.3.2. *Cointegration test using the Johansen methodology*

The paper also tested for cointegration relationships between markets during the two sub-periods using the Johansen methodology. This test is based on the relationship between the rank of the matrix and its characteristic roots. The starting point from the autoregressive vector model (VAR) of the order p is presented as follows:

$$Y_t = A_1 Y_{t-1} + \ldots A_p Y_{t-p} + BX_t + \varepsilon_t. \quad (1)$$

3 The Engle and Granger method is valid for the co-integrated series of Order 1.

Table 3. Bivariate cointegration between the U.S. and other countries: stability vs. crisis periods

Bivariate Cointegration			
American region	**ADF**	**Critical value 5%**	**Decision**
U.S.–CANADA	–2.801*	–1.95	Yes
	(–3.5057)*	(–1.95)	(Yes)
U.S.–MEXICO	–2.4275*	–1.95	Yes
	(–1.8037)	(–1.95)	(No)
U.S.–ARGENTINA	–3.64*	–1.95	Yes
	(–1.9121)	(–1.95)	(No)
U.S.–BRAZIL	–2.2971*	–1.95	Yes
	(–2.1953)	(–1.95)	(Yes)
CANADA–U.S.	–2.591*	–1.95	Yes
	(–3.5401)*	(–1.95)	(Yes)
MEXICO–U.S.	–1.925	–1.95	No
	(–2.7078)	(–3.41)	(No)
ARGENTINA–U.S.	–3.626*	–1.95	Yes
	(–4.2095)*	(–3.41)	(Yes)
BRAZIL–U.S.	–1.9283	–1.95	No
	(–2.1028)*	(–1.95)	(Yes)
Asian region	**ADF**	**Critical value 5%**	**Decision**
U.S.–JAPAN	–2.5157*	–1.95	Yes
	(–1.5695)	(–1.95)	(No)
U.S.–HONG KONG	–3.0712*	–1.95	Yes
	(–2.6233)*	(–1.95)	(Yes)
U.S.–INDIA	–2.8729*	–1.95	Yes
	(–2.3529)*	(–1.95)	(Yes)
U.S.–INDONESIA	–2.5761*	–1.95	Yes
	(–2.1557)*	(–1.95)	(Yes)
U.S.–MALAYSIA	–2.0115*	–1.95	Yes
	(–2.2115)*	(–1.95)	(Yes)
U.S.–KOREA	–1.9512*	–1.95	Yes
	(–2.3546)*	(–1.95)	(Yes)
U.S.–CHINA	–1.6992	–1.95	No
	(–1.7826)	(–1.95)	(No)
U.S.–SINGAPORE	–3.1341*	–1.95	Yes
	(–3.1302)*	(–1.95)	(Yes)
JAPAN–U.S.	–2.4492*	–1.95	Yes
	(–2.0484)*	(–1.95)	(Yes)
HONG KONG–U.S.	–2.7446*	–1.95	Yes
	(–2.6366)*	(–1.95)	(Yes)
INDIA–U.S.	–2.4975*	–1.95	Yes
	(–2.2424)*	(–1.95)	Yes
AUSTRALIA–U.S.	–2.7291*	–1.95	Yes
	(–2.0905)*	(–1.95)	(Yes)
INDONESIA–U.S.	–2.0866*	–1.95	Yes
	(–3.3271)	(–3.41)	(No)
MALAYSIA–U.S.	–1.4662	–1.95	No
	(–2.8798)	(–3.41)	(No)
KOREA–U.S.	–1.2885	–1.95	No
	(–3.1042)	(–3.41)	(No)
CHINA–U.S.	1.3667	–1.95	No
	(–1.6741)	(–1.95)	(No)
SINGAPORE–U.S.	–2.8228*	–1.95	Yes
	(–3.0257)*	(–1.95)	(Yes)

Note: * is statistical significance at the 5% level.

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where Y_t – A vector with K endogenous variable, X_t – A vector with N exogenous variables, $A_1,\ldots,A_p$ and B – matrices of the coefficients to be estimated, ε_t – A vector of innovations not correlated with the endogenous variables.

The number of delays, p, is determined using the Schwarz Bayesian information criterion (SBIC). The optimal number of delays retained is equal to 1, both for stability and crisis periods.

Subsequently, we proceed to Johansen's multivariate cointegration test between the U.S. stock market and the other markets examined from the two sub-periods: stability and crisis.

To test the presence of cointegration relationships between markets ($H_{0:}$ X_t is cointegrated of rank *r*), Johansen (1988) recommends two different tests, i.e., the trace test and the maximum eigenvalue test.

$$J_{trace} = -T \sum_{i=r+1}^{n} \ln\left(1-\hat{\lambda}_i\right), \qquad (2)$$

$$J_{\max} = -T \ln\left(1-\hat{\lambda}_{r+1}\right), \qquad (3)$$

where T – the sample size, λ_i – the ith eigenvalue.

Table 4 reports the results from Johansen's cointegration tests between the U.S. stock market and the other markets examined during the stability and crisis periods. The trace and maximum eigenvalue statistics are used to test for multivariate cointegration. Table 3 shows the absence of a cointegration relationship between the U.S. stock market and other markets of the American region during both studied periods using trace tests.

For the Asian region, we start first with the Johansen cointegration results during the stability period. The findings show that there are, at most, two ($r <= 2^*$) cointegration relationships using trace statistics, while the use of the maximum eigenvalue statistic indicates a cointegration relationship at the 1%, 5%, and 10% thresholds.

The crisis period shows that there is, at most ($r <= 1^*$), a cointegrating relationship between the U.S. stock market and the Asian stock markets at the 1%, 5%, and 10% thresholds, using the trace and maximum eigenvalue statistics (72.00> 62.42 and 57.61> 57.00).

The paper first estimated both the Engel-Granger and Johansen cointegration tests. Then, examined the asymmetric cointegration relationship between variables using Enders and Siklos's (2001) threshold autoregressive (TAR) model. It failed to detect asymmetric cointegration relationships among variables.

Table 4. Multivariate cointegration test for the American and Asian regions: stability vs. crisis periods

Panel A. American region					
Trace test					
H0	**Eigenvalues**	λ_{trace}	**Critical value 10%**	**Critical value 5%**	**Critical value 1%**
r=0	1.717e–02	51.94	66.49	70.60	78.87
	(0.022)	(64.14)	(66.49)	(70.60)	(78.87)
r <= 1	1.092e–02	31.29	45.23	48.28	55.43
	(0.018)	(38.32)	(45.23)	(48.28)	(55.43)
r <= 2	8.608e–03	18.19	28.71	31.52	37.22
	(0.009)	(16.73)	(28.71)	(31.52)	(37.22)
r <= 3	6.559e–03	7.88	15.66	17.95	23.52
	(0.003)	(5.38)	(15.66)	(17.95)	(23.52)
r <= 4	3.153e–05	0.04	6.50	8.18	11.65
	(0.0007)	(0.92)	(6.50)	(8.18)	(11.65)
Maximum eigenvalue test					
H0	**Eigenvalues**	λ_{max}	**Critical value 10%**	**Critical value 5%**	**Critical value 1%**
r=0	1.717e-02	20.65	30.84	33.32	38.78
	(0.022)	(25.82)	(30.84)	(33.32)	(38.78)
r <= 1	1.092e-02	13.10	24.78	27.14	32.14
	(0.0186)	(21.59)	(24.78)	(27.14)	(32.14)

Table 4 (cont.). Multivariate cointegration test for the American and Asian regions: stability vs. crisis periods

Maximum eigenvalue test					
H0	**Eigenvalues**	λ_{max}	**Critical value 10%**	**Critical value 5%**	**Critical value 1%**
r <= 2	8.608e-03	10.31	18.90	21.07	25.75
	(0.009)	(11.35)	(18.90)	(21.07)	(25.75)
r <= 3	6.559e-03	7.85	12.91	14.90	19.19
	(0.003)	(4.46)	(12.91)	(14.90)	(19.19)
r <= 4	3.153e-05	0.04	6.50	8.18	11.65
	(0.0007)	(0.92)	(6.50)	(8.18)	(11.65)
Panel B. Asian region					
Trace test					
H0	**Eigenvalues**	λ_{trace}	**Critical value 10%**	**Critical value 5%**	**Critical Value 1%**
r=0*	0.066	295.78	226.34	232.49	246.27
	(0.060)	(279.91)	(226.34)	(232.49)	(246.27)
r <= 1*	0.040	213.68	186.54	192.84	204.79
	(0.048)	(207.9)	(186.54)	(192.84)	(204.79)
r <= 2*	0.038	164.03	151.38	157.11	168.92
	(0.041)	(150.30)	(151.38)	(157.11)	(168.92)
r <= 3	0.033	117.73	118.99	124.25	136.06
	(0.030)	(102.19)	(118.99)	(124.25)	(136.06)
r <= 4	0.021	77.53	90.39	85.18	104.20
	(0.020)	(24.25)	(36.25)	(39.43)	(44.59)
r <= 5	0.015	52.00	66.49	70.60	78.87
	(0.011)	(13.43)	(30.84)	(33.32)	(38.78)
r <= 6	0.014	33.21	45.23	48.28	55.43
	(0.010)	(28.99)	(45.23)	(48.28)	(55.43)
r <= 7	0.008	16.20	28.71	31.52	37.22
	(0.009)	(17.36)	(28.71)	(31.52)	(37.22)
r <= 8	0.004	5.85	15.66	17.95	23.52
	(0.004)	(6.21)	(15.66)	(17.95)	(23.52)
r <= 9	0.0007	0.90	6.50	8.18	11.65
	(0.0008)	(0.98)	(6.50)	(8.18)	(11.65)
Maximum eigenvalue test					
H0	**Eigenvalues**	λ_{max}	**Critical value 10%**	**Critical Value 5%**	**Critical Value 1%**
r=0*	0.066	82.10	59.00	62.42	68.61
	(0.060)	(72.00)	(59.00)	(62.42)	(68.61)
r <= 1	0.040	49.66	54.01	57.00	63.37
(r <= 1*)	(0.048)	(57.61)	(54.01)	(57.00)	(63.37)
r <= 2	0.038	46.30	48.43	51.07	57.07
	(0.041)	(48.11)	(48.43)	(51.07)	(57.07)
r <= 3	0.033	40.20	42.06	44.91	51.30
	(0.030)	(35.52)	(42.06)	(44.91)	(51.30)
r <= 4	0.021	25.52	36.25	39.43	44.59
	(0.020)	(24.25)	(36.25)	(39.43)	(44.59)
r <= 5	0.015	18.79	30.84	33.32	38.78
	(0.011)	(13.43)	(30.84)	(33.32)	(38.78)
r <= 6	0.014	17.01	24.78	27.14	32.14
	(0.010)	(11.62)	(24.78)	(27.14)	(32.14)
r <= 7	0.008	10.35	18.90	21.07	25.75
	(0.009)	(11.15)	(18.90)	(21.07)	(25.75)
r <= 8	0.0041	4.95	12.91	14.90	19.19
	(0.004)	(5.24)	(12.91)	(14.90)	(19.19)
r <= 9	0.0007	0.90	6.50	8.18	11.65
	(0.0008)	(0.98)	(6.50)	(8.18)	(11.65)

Notes: The decision rule for this test is to reject the null hypothesis of the r relation of cointegration when the TR statistic is greater than its critical value.

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3.4. Estimation of the VAR and VECM models and results from causality tests

The next step is to test the causality of Granger for the different returns of stock indexes considered.

Table 5 shows the linear causal relationships that exist between the US market and other studiedamarkets for both sub-periods. In the stability period, 8 unidirectional relationships (USA to Argentina, Mexico to USA, USA to Malaysia, USA to Korea, USA to Japan, USA to Hong Kong, USA to China, and USA to Singapore) are distinguished and 3 other bidirectional (USA and Brazil, USA and India, USA and Indonesia). During the period of the subprime crisis, the number of two-way relationships decreases to 2 causal relationships only, between the USA and Canada and between the USA and Hong Kong. This table also shows the results of the existence or absence of a causal relationship between the different countries studied.

Table 5. Linear Granger-causality tests between the US and other markets – stability period (crisis period)

Linear causality test			
American region	**Fisher**	**P-value**	**Decision**
U.S.–CANADA	8.496***	0.0036	Yes
	(14.392***)	(0.00015)	(Yes)
U.S.–MEXICO	2.565	0.1094	No
	(0.0872)	(0.767)	(No)
U.S.–ARGENTINA	3.374*	0.066	Yes
	(0.598)	(0.439)	(No)
U.S.–BRAZIL	3.963**	0.046	Yes
	(2.142)	(0.143)	(No)
CANADA–U.S.	1.192	0.274	No
	(8.80***)	(0.0030)	(Yes)
MEXICO–U.S.	4.541**	0.033	Yes
	(3.8671**)	(0.049)	(Yes)
ARGENTINA–U.S.	0.002	0.962	No
	(0.134)	(0.714)	(No)
BRAZIL–U.S.	4.528**	0.033	Yes
	(0.750)	(0.386)	(No)
Asian region	**Fisher**	**P-value**	**Decision**
U.S.–JAPAN	13.025***	0.0003	Yes
	(4.625**)	(0.031)	(Yes)
U.S.–HONG KONG	24.035***	1.076e–06	Yes
	(4.058**)	(0.044)	(Yes)
U.S.–INDIA	8.8266**	0.0030	Yes
	(0.203)	(0.652)	(No)
U.S.–INDONESIA	11.945	0.0005	Yes
	(0.194)	(0.659)	(No)
Asian region	**Fisher**	**P-value**	**Decision**
U.S.–MALAYSIA	11.514***	0.0007	Yes
	(0.0521)	(0.8193)	(No)
U.S.–KOREA	12.07***	0.0005	Yes
	(0.366)	(0.5450)	(No)
U.S.–CHINA	3.768*	0.052	Yes
	(0.913)	(0.339)	(No)
U.S.–SINGAPORE	24.63***	7.955e–07	Yes
	(0.277)	(5.985e–01)	(No)
JAPAN–U.S.	0.493	0.4826	No
	(0.005)	(0.942)	(No)
HONG KONG–U.S.	1.576	2.095e–01	No
	(3.667*)	(0.055)	(Yes)
INDIA–U.S.	3.215*	0.073	Yes
	(3.304*)	(0.069)	(Yes)
INDONESIA–U.S.	2.741*	0.098	Yes
	(3.475*)	(0.0625)	(Yes)
MALAYSIA–U.S.	0.188	0.664	No
	(4.921**)	(0.026)	(Yes)
KOREA–U.S.	0.138	0.7098	No
	(5.031**)	(0.025)	(Yes)
CHINA–U.S.	0.034	0.852	No
	(2.842*)	(0.092)	(Yes)
SINGAPORE–U.S.	1.803	1.79e–01	No
	(15.954***)	(6.902e–05)	(Yes)

Note: *, **, *** are significance levels at 10%, 5% and 1%, respectively.

The paper further tests for Granger's sense of causality of different returns from the stock indices examined. The causality tests were applied based on the following equations:

$$\Delta x_t = \alpha_x + \sum_{i=1}^{k} \beta_{x,i} \Delta x_{t-i} + \sum_{i=1}^{k} \tau_{x,i} \Delta y_{t-i} + \varepsilon_{x,t}, \quad (4)$$

$$\Delta y_t = \alpha_y + \sum_{i=1}^{k} \beta_{y,i} \Delta y_{t-i} + \sum_{i=1}^{k} \tau_{y,i} \Delta x_{t-i} + \varepsilon_{y,t}. \quad (5)$$

In case of cointegration, an error correction term (ECT) is integrated into the differentiated VAR equation, and the order VAR (k) then becomes a VECM of order (K-1). The paper uses, in this case, the following equations:

$$\Delta x_t = \alpha_x + \sum_{i=1}^{k} \beta_{x,i} \Delta x_{t-i} + \sum_{i=1}^{k} \tau_{x,i} \Delta y_{t-i} + \delta_x ECT_{x,(t-1)} + \varepsilon_{x,t}, \quad (6)$$

$$\Delta y_t = \alpha_y + \sum_{i=1}^{k} \beta_{y,i} \Delta y_{t-i} + \sum_{i=1}^{k} \tau_{y,i} \Delta x_{t-i} + \delta_y ECT_{y,(t-1)} + \varepsilon_{y,t}. \quad (7)$$

At this level, we distinguish between two types of causality (Sander & Kleimer, 2003):

- The non-causality in the short term that is tested by hypothesis H_0: $\tau_{x,i} = 0$. If H_0 is rejected, then y causes x in the short term in the Granger sense.
- The non-causality in the long term that is tested by hypothesis H_0: $\delta_x = 0$. Similarly, if H_0 is rejected, then y causes x in the long term in the Granger sense.

The hypotheses tested are then:

$$\begin{cases} H_0 : \tau_{j,i} = 0 : non-causality\ in\ the\ short\ term \\ H_1 : \tau_{j,i} \neq 0 :\ causality\ in\ the\ short\ term \end{cases}, \quad (8)$$

$$\begin{cases} H_0 : \delta = 0 : non-causality\ in\ the\ long\ term \\ H_1 : \delta \neq 0 :\ causality\ in\ the\ long\ term \end{cases}. \quad (9)$$

The parameters $\tau_{j,i}$ and δ indicate evidence of causality in the short and long terms.

Table 6 summarizes the causality test results during the stability period and the subprime crisis period for the Asian and American regions compared with the U.S. Table 6 shows that the number of short-term causal relationships between the U.S. and the other studied countries is 13 during a period of stability and 14 in times of crisis. Short-term causal relationships are found between U.S.-Mexico / Canada-U.S. / Japan- U.S. / Hong Kong-U.S. during both stability and crisis periods. On the other hand, the long term illustrates 10 causal relationships during the stability period and nine during times of crisis. Japan-U.S. and Hong Kong-U.S. present long-term causal relationships during stability and crisis periods.

The paper uses the Granger causality approach to identify cases of pure contagion and financial interconnection between the studied markets. Indeed, the existence of short-term or long-term causal relationships between the stock markets is evidence of pure contagion. In addition, if this relationship still exists during a period of stability, it indicates a transmission of crisis from one market to another through an interconnection between the originating crisis market and the country affected by the crisis (Marais & Bates, 2006).

Table 7 identifies cases of pure contagion and financial interconnection between the U.S. market and the other studied markets, respectively. The study illustrated eight cases of pure contagion (e.g., U.S.-Japan, U.S.-Hong Kong, U.S.-Malaysia, U.S.-Singapore, China-U.S.) and 10 cases of financial interconnection between the U.S. market and the other studied markets (e.g., U.S.-Mexico,

Table 6. Results from the short- and long-term causality test for the region: U.S. with other countries – period of stability (crisis period)

	Short term			Long term		
American region	$\tau_{j,i}$	**Student's t**	**Probability**	δ	**Student's t**	**Probability**
U.S.–CANADA	0.0032	0.783	0.4340	–0.007	–1.152	0.2494
	(–0.0058)	(–1.353)	(0.1761)	(–0.0084)	(–0.996)	(0.3194)
U.S.–MEXICO	0.0032	2.112	0.0348 *	–0.0081	–1.886	0.0595*
	(0.0956)	(1.967)	(0.049 *)	(–)	(–)	(–)
U.S.–ARGENTINA	–0.006	–0.492	0.622	–0.0080	–1.269	0.2048
	(–0.009)	(–0.366)	(0.714)	(–)	(–)	(–)
U.S.–BRAZIL	0.0008	1.287	0.198	–0.0065	–1.175	0.240
	(0.0001)	(0.173)	(0.863)	(–0.0032)	(–1.245)	(0.213)
CANADA–U.S.	1.080	3.521	0.0004 ***	–0.0065	–1.776	0.0759
	(0.9605)	(2.384)	(0.017 *)	(–0.0158)	(–1.681)	(0.093)
MEXICO–U.S.	0.080	1.602	0.109	–	–	–
	(–0.0126)	(–0.295)	(0.768)	(–)	(–)	(–)
ARGENTINA–U.S.	0.146	1.831	0.067	–0.0149	–2.681	0.007 ***
	(–0.068)	(–0.699)	(0.485)	(0.0024)	(1.259)	(0.208)
BRAZIL–U.S.	0.143	1.991	0.0467 *	–	–	–
	(1.6009)	(0.644)	(0.5197)	(–0.0040)	(–0.941)	(0.347)

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Table 6 (cont.). Results from the short- and long-term causality test for the region: U.S. with other countries – period of stability (crisis period)

	Short term			Long term		
American region	$\tau_{j,i}$	**Student's t**	**Probability**	δ	**Student's t**	**Probability**
U.S.–JAPAN	–0.0007	–0.441	0.6592	–0.0025	–0.602	0.5476
	(–0.0467)	(–1.796)	(0.0728*)	(–)	(–)	(–)
U.S.–HONG KONG	–0.002	–1.107	0.2685	–0.007	–1.156	0.2480
	(–0.0005)	(–0.430)	(0.6675)	(–0.0096)	(–2.097)	(0.0362 **)
U.S.–INDIA	–0.0014	–0.681	0.4959	–0.0079	–1.713	0.0870
	(–0.0001)	(–0.097)	(0.9230)	(–0.0074)	(–2.250)	(0.0247 **)
U.S.–INDONESIA	–0.0366	–1.970	0.0491 **	–0.006	–1.329	0.1841
	(0.006)	(0.465)	(0.6423)	(–0.0078)	(–2.467)	(0.0138 **)
U.S.–MALAYSIA	–0.0039	–0.100	0.9205	–0.003	–0.835	0.4037
	(–0.0612)	(–1.534)	(0.125)	(–0.0105)	(–2.816)	(0.0049 ***)
U.S.–KOREA	0.028	1.418	0.1563	–0.0031	–0.903	0.36655
	(0.0008)	(0.042)	(0.966)	(–0.0105)	(–2.649)	(0.008 ***)
U.S.–CHINA	–0.002	–0.189	0.8499	–	–	–
	(0.016)	(0.809)	(0.419)	(–)	(–)	(–)
U.S.–SINGAPORE	–0.0119	–0.940	0.3474	–0.0077	–1.193	0.2330
	(0.0345)	(2.697)	(0.007 ***)	(–0.0258)	(–3.874)	(0.001***)
JAPAN–U.S.	6.707	15.103	<2e–16 ***	–0.0081	–2.387	0.0171 **
	(4.607)	(14.721)	(< 2e–16 ***)	(–0.0067)	(–2.300)	(0.0216 **)
HONG KONG–U.S.	6.789	15.927	< 2e–16 ***	–1.1e–02	–2.824	0.0048 ***
	(7.803)	(12.218)	(< 2e–16 ***)	(0.120)	(–1.556)	(0.0079***)
INDIA–U.S.	3.822	10.025	<2e–16 ***	–5.38e–03	–2.266	0.0237 **
	(2.712)	(6.114)	(1.33e–09 ***)	(–0.0032)	(–1.016)	(0.310)
INDONESIA–U.S.	0.4885	11.300	< 2e–16 ***	–0.0051	–2.903	0.0037 ***
	(0.229)	(7.880)	(7.55e–15 ***)	(–)	(–)	(–)
MALAYSIA–U.S.	0.263	11.189	< 2e–16 ***	–	–	–
	(0.145)	(7.729)	(2.36e–14 ***)	(–)	(–)	(–)
KOREA–U.S.	0.570	12.418	<2e–16 ***	–	–	–
	(0.2628)	(8.624)	(< 2e–16 ***)	(–)	(–)	(–)
CHINA–U.S.	–0.0355	–0.659	0.5103	–	–	–
	(0.2423)	(5.927)	(4.08e–09 ***)	(–)	(–)	(–)
SINGAPORE–U.S.	0.9647	15.500	< 2e–16 ***	–0.0106	–3.092	0.0020***
	(0.5809)	(8.320)	(2.48e–16 ***)	(0.0014)	(0.258)	(0.797)

Note: *, **, and *** are significance levels at 10%, 5% and 1%, respectively.

Japan-U.S., Malaysia-U.S., and Singapore-U.S.). Moreover, high co-movements in the short term (6 short-term causalities) and low co-movements in the long term (1 long-term causality) for cases of financial interconnection are found.

These results are consistent with Davidson (2020) who shows that during the global financial crisis, Mexico was contagious due to existing interconnections with the U.S. However, the results are inconsistent with Dewandaru et al. (2016) who argue that the subprime crisis had fundamentals-based contagion. Instead, they find high co-movements in the long term and low co-movements in the short term.

This paper makes several contributions to the literature. First, it identifies the beginning date of the subprime crisis using a novel technique. The research extends Boyer et al. (2006), Rodriguez (2007), Baur (2012), and Dimitriou et al. (2013), who used the Markov switching dynamic regression (MS-DR) model to detect the crisis date. It uses the wavelet technique to distinguish between stability and crisis periods. Second, to the best of our knowledge, no study has investigated the existence of pure contagion and financial interconnection between the U.S. and other American and Asian regions in the subprime crisis context. This study distinguishes between cases of pure contagion and

Table 7. Pure contagion and financial interconnection identification: U.S. with other countries

Cases of pure contagion		Cases of interconnection	
Causal relationships	Type of causality	Causal relationships	Type of causality
U.S.-JAPAN	Short term	U.S.-MEXICO	Short term
U.S.-HONG KONG	Long term	CANADA-U.S.	Short and long term
U.S.-INDONESIA	Long term	U.S.-INDIA	Long term
U.S.- MALAYSIA	Long term	JAPAN-U.S.	Short and long term
U.S.-KOREA	Long term	HONG KONG-U.S.	Short and long term
U.S.-SINGAPORE	Short and long term	INDIA-U.S.	Short term
CHINA-U.S.	Short term	INDONESIA-U.S.	Short term
		MALAYSIA-U.S.	Short term
		KOREA-U.S.	Short term
		SINGAPORE–U.S.	Short term

Note: The existence of short-term or long-term causal relationships between the stock markets is evidence of pure contagion. If this relationship still exists during a period of stability, it indicates an interconnection between the originating crisis market and the country affected by the crisis.

cases of interconnection and the type of causality. Finally, unlike studies that have used the cointegration technique in the 1997 Asian crisis context – such as Masih and Masih (1999), Tan (1998), and Yang et al. (2005) – it adopts cointegration and Granger causality approaches by estimating vector autoregressive (VAR) and vector error correction models (VECM).

CONCLUSION

This study examines the existence of pure contagion and financial interconnections in the subprime crisis context. We adopt the cointegration and Granger causality approach by estimating the VAR and VECM models. Based on the daily series from stock indices in American and Asian countries, the phenomenon of pure contagion and financial interconnection is investigated.

The paper puts forward eight cases of pure contagion and ten cases of financial interconnection by studying markets in the U.S. and other American countries and markets from the Asian region. In addition, there are high co-movements in the short term and low co-movement in the long term in cases of financial interconnection. The results show that the subprime crisis has affected the Asian region (Japan, Hong Kong, Indonesia, Malaysia, etc.). The findings support the contagious nature of the subprime crisis between the U.S. and Asian countries and the U.S. and other American countries.

The results have several implications for investors who seek to diversify their portfolios internationally. In terms of portfolio diversification, when equity returns are cointegrated, we note that in the long run, these stocks have high long run correlations and are therefore unnecessary redundant diversifiers in portfolios. Then, from an investment standpoint, there is less potential gain from international portfolio diversification. Furthermore, the existence of a short-term causal relationship between the U.S. and other markets helps investors make investment decisions.

The results also have strong implications for policymakers. Indeed, financial contagion between international stock markets can help decision-makers develop the existing financial system and make it more immune to the transmission of crises. Moreover, the study of such phenomena is important to better understand the effectiveness of international financial institutions' actions and policies in the context of crises.

Future research avenues may focus on the effects of contagion and financial interconnections on the real economy. It suggests replicating the approach used in this study in the context of the COVID-19 pandemic and comparing it with other approaches.

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