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Emerging markets and financial crises: Regional, global or isolated shocks?

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ABSTRACT

This paper investigates financial contagion of three emerging market crises of the late 1990s, as well as the subprime crisis of 2007, focusing on financial markets of emerging economies, USA and 2 global indices. Conventional cointegration and vector error correction analysis show long and short run dynamics only among emerging stock markets during the Russian and the Asian crises, for both stock and bond markets during the subprime crisis, while the Argentine turmoil has no impact on any of the examined markets. Further analysis into a multivariate time-varying asymmetric framework provides evidence on the global impact of the Russian default, the contagion effects of the subprime crisis, the regional aspect of the Asian crisis and the isolated nature of the Argentine turmoil. Moreover, stock markets seem to constitute a stronger transmission mechanism during the three contagious crises. Our findings have crucial implications for international investors, policy makers and multi-lateral organizations.

1. Introduction

A central issue in asset allocation and risk management is whether financial markets become more interdependent during financial crises. This issue has acquired great importance among academics and practitioners, especially since the appearance of several emerging market crises of the 1990s

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(Mexican currency crisis in 1994–1995, Asian crisis in 1997, Russian default in 1998, Argentine crisis in 1999–2001, and Brazilian stock market crash in 1997–1998). Until then, financial crises models were developed with regard to crises as events occurring in individual countries. However, those crises episodes refocused the empirical research on the examination of contagion effects and the inter-regional or inter-continental nature of the shocks. Common to the majority of these events was the fact that the turmoil originated in one market extended to a wide range of markets and countries, in a way that was hard to explain on the basis of changes in fundamentals (Rodriquez, 2007).

There is an extensive literature on financial contagion during several crises of the 1980s and 1990s (see Dornbusch et al., 2000; Kaminsky et al., 2003, for excellent surveys). To measure volatility spillovers and contagion among markets, early research used a range of different methodologies, such as the principle components model (e.g., Calvo and Reinhart, 1996), spillover models (e.g., Glick and Rose, 1999), cointegration and vector error correction models (e.g., Sheng and Tu, 2000), models of asymmetries and non-linearities (e.g., Baur, 2003), and models of interdependence (e.g., Bekaert et al., 2005). Furthermore, the existence of financial contagion has been studied mainly around the notion of “correlation breakdown” (a statistically significant increase in correlation during the crash period) (e.g., King and Wadhwani, 1990; Calvo and Reinhart, 1996). However, since the thought-provoking paper by Forbes and Rigobon (2002), a number of limitations to the literature on financial contagion have been highlighted (e.g., a heteroskedasticity problem when measuring correlations, a problem with omitted variables and the need for a dynamic increment in the regressions, affecting at least in the second moments correlations and covariances). To overcome those restrictions and provide sufficient evidence of contagion, researchers have been already using more sophisticated approaches, such as models of conditional asymmetries and correlations (e.g., Chiang et al., 2007; Kenourgios et al., 2011), regime-switching models (e.g., Pelletier, 2006), and dynamic copulas with or without regimes (e.g., Rodriquez, 2007; Okimoto, 2008).

This paper investigates the existence of a correlated-information channel, through which contagion can be viewed as the transmission of information from more-liquid markets or markets with more rapid price discovery to other markets, focusing initially on three major emerging market crises (Asian crisis, Russian default and Argentine turmoil).[1] The analysis covers both equity and bond markets in selected emerging market economies (EMEs) of various regions (Latin America, Asia, Europe, Middle East and Africa), as well as USA and 2 global indices for equities and bonds, for comparative reasons. To expand the scope and the contribution of our research, we also investigate the contagion effects of the subprime crisis of 2007–2008 in the examined EMEs.[2] Our purpose is to identify in a broader framework the propagation mechanism of crises with different characteristics occurred in emerging (Asian currency crisis, Russian and Argentinean government defaults) and developed countries/regions (U.S. subprime crisis), and elucidate how vulnerable emerging financial markets are to both emerging and global shocks. To serve this purpose, we maintain, following similar studies in the literature (e.g., Forbes and Rigobon, 2002; Bekaert et al., 2005), an equivalently strict definition of contagion as the increase in the probability of crisis, beyond the linkages in fundamentals, and the rapid increase in co-movements among markets during a crisis episode. Understanding the nature

[1] The contagion literature identifies at least three possible contagion mechanisms: (i) a correlated-information channel (Kaminsky et al., 2003; King and Wadhwani, 1990, among others); (ii) a liquidity channel, through which contagion occurs through a liquidity shock across all markets (e.g., Allen and Gale, 2000, among others); (iii) a risk-premium channel, through which contagion occurs as negative returns in the distressed market affect subsequent returns in other markets via a time-varying risk premium (e.g., Acharya and Pedersen, 2005, and others). In this paper, we restrict our investigation only on the first contagion mechanism, due to the lack of availability of consistent and compatible financial data in emerging markets.

[2] Asian crisis contagion clearly receives the highest share of attention in the literature (e.g., Glick and Rose, 1999; Baig and Goldfajn, 1999; Sheng and Tu, 2000; Chiang et al., 2007; Kenourgios et al., 2011). On the other hand, little empirical investigation of the Russian default has been performed, while there is limited consensus regarding its contagious effects (for example, Gelos and Sahay, 2000, find no contagion, while Forbes, 2000, and Dungey et al., 2007, confirm the contagion effect). On the contrary, empirical evidence on the contagion of the Argentinean default in global financial markets is surprisingly scarce (e.g., Boschi, 2005). Although the literature on the international impact of the U.S. subprime crisis is still developing, only few studies focus in EMEs. For example, Dooley and Hutchison (2009) provide evidence on the decoupling of emerging markets from early 2007 to summer 2008, but after that point confirms their recoupling due to the deteriorating situation in the U.S. financial system and real economy, while Aloui et al. (2011) find strong evidence of time-varying dependence between each of the BRIC equity markets and the U.S. markets.

and the differences in crises dynamics has crucial implications for international investors, portfolio managers, policy makers and multi-lateral organizations.

Initially, we employ a conventional empirical analysis (Johansen cointegration tests and vector error correction model), and find significant long and short-run market dynamics only among emerging stock markets during the Russian and the Asian crises, for both stock and bond markets during the subprime crisis, while the Argentine crisis has no impact on any of the examined financial markets. However, to overcome the limitations of the contagion literature, we also apply a recently developed GARCH process, the asymmetric generalized dynamic conditional correlation (AG-DCC) model developed by Cappiello et al. (2006), who generalized the DCC-GARCH model of Engle (2002). Average correlations between the crisis country and all other countries during stable and crisis periods are estimated, which allow capturing conditional asymmetries in both volatilities and correlations in a multivariate framework and determining whether cross-market correlation dynamics (contagious effects) are driven by changes in macroeconomic fundamentals or by behavioral reasons. Results provide evidence on the global contagion effect of the subprime crisis and the Russian default, the intra-regional aspect of the Asian crisis and the decoupling of the examined national and global market indices to the Argentine turmoil.

This paper contributes to the existing literature in the following aspects. First, we provide new evidence on financial theory of contagion by examining the existence of an asymmetric propagation mechanism for financial crises which have different characteristics and origins. The AG-DCC GARCH model applied in this paper is well suited to examine asymmetric responses to negative shocks (stronger contagion during negative shocks) and, to the best of our knowledge, has not been used before to test the contagion hypothesis for all four financial crises. Second, we differentiate our analysis from previous studies, focusing on emerging stock and bond markets from various regions around the world, instead of individual mature economies, given the limited research on EMEs. In this framework, we identify which of the two asset markets are more prone to financial contagion in EMEs. The propagation of contagion into bond markets of emerging economies with heavily exposure on debt as primary financing source (IMF International Capital Markets, September 1998, and IMF Global Financial Stability Report, September 2006) constitutes a topic of little empirical investigation, since the majority of existing studies focus mainly on the transmission of shocks across foreign exchange and stock markets. Third, the analysis of contagion of the subprime crisis is also of great importance, given the existing debate on whether the contagion effect on EMEs has been muted and uneven (decoupling hypothesis) or not. Fourth, examining differences in crises dynamics reveals several explanatory factors and contributes to the debates regarding the resilience and sustainability of emerging-market policy performance and the construction of a new international (or regional) financial architecture.

The structure of the paper is organized as follows. Section 2 presents the methodologies applied to examine financial contagion. The data used for the empirical analysis is presented in Section 3. Section 4 reports the empirical results. Finally, concluding remarks are stated in Section 5.

2. Methodology

2.1. Cointegration tests

We test for the presence and number of cointegrating relationships among the emerging markets on a vector error correction model, applying the procedure advanced by Johansen (Johansen, 1991; Johansen and Juselius, 1990).

Defining a vector z_t of n endogenous variables, it is possible to specify the following data generating process and model z_t as an unrestricted vector autoregression (VAR) involving up to k– lags of z_t:

$$Z_t = A_1 z_{t-1} + A_2 z_{t-2} + \ldots + A_\lambda z_{t-k} + u_t \qquad u_t \sim IN(0, \sum) \tag{1}$$

where z_t is a $(n \times 1)$ matrix, and each of A_i is a $(n \times n)$ matrix of parameters. Then Eq. (1) can be reformulated into a VECM form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + u_t \quad \text{or}$$

$$\Delta z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \Pi z_{t-k} + u_t \tag{2}$$

where $\Gamma_i = -(I - A_1 - \ldots - A_i)$, $(i = 1, \ldots, k-1)$, Γ_i are interim multipliers, and $\Pi = -(\mathrm{I} - A_1 - \ldots - A_k)$. Testing for cointegration is related to the consideration of the rank of Π, that is finding the number of r linearly independent in Π. The trace and maximum eigenvalue tests are used to identify the number of cointegrating vectors among variables. The presence of cointegrating vectors supports the application of a dynamic VECM that depicts the feedback process and adjustment speed of short-run deviations towards the long-run equilibrium and reveals short-run dynamics in any market relative to the other markets.

2.2. AG-DCC model

Volatilities and correlations measured from historical data may miss changes in risk. Thus, Cappiello et al. (2006) investigate properties of international equity returns generalizing the DCC-GARCH model of Engle (2002). This process interprets asymmetries broader than just within the class of GARCH models (does not assume constant correlation coefficients over the sample period), allows for series-specific news impact and smoothing parameters, permits conditional asymmetries in correlation dynamics and accounts for heteroskedasticity directly by estimating correlation coefficients using standardized residuals. Moreover, this specification overcomes the problem with omitted variables and is well suited to investigate the presence of asymmetric responses in conditional variances and correlations during periods of negative shocks.

In this paper, we estimate univariate volatility using a GARCH (1,1) model (Bollerslev, 1986), and the standardized residuals, $\varepsilon_{i,t} = r_{i,t}/\sqrt{h_{i,t}}$, are used to estimate the correlation parameters. The evolution of the correlation in the standard DCC model (Engle, 2002) is given by:

$$Q_t = (1 - a - b)\bar{P} + a\varepsilon_{t-1}\varepsilon'_{t-1} + bQ_{t-1} \tag{3}$$

$$P_t = Q_t^{*-1} Q_t Q_t^{*-1} \tag{4}$$

where, $\bar{P} = E[\varepsilon_t \varepsilon'_t]$ and α and b are scalars such that $\alpha + b < 1$. $Q_t^* = \left[q_{iit}^*\right] = [\sqrt{q_{iit}}]$ is a diagonal matrix with the square root of the ith diagonal element of Q_t on its ith diagonal position. As long as Q_t is positive definite, Q_t^* is a matrix which guarantees $P_t = Q_t^{*-1} Q_t Q_t^{*-1}$ is a correlation matrix with ones on the diagonal and every other element ≤ 1 in absolute value. The model described by Eqs. (3) and (4), however, does not allow for asset-specific news and smoothing parameters or asymmetries.

Cappiello et al. (2006) modify the correlation evolution equation as:

$$Q_t = (\bar{P} - A'\bar{P}A - B'\bar{P}B - G'\bar{N}G) + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + G'n_{t-1}n'_{t-1}G + B'Q_{t-1}B) \tag{5}$$

where, A, B and G are $k \times k$ parameter matrices, $n_t = I[\varepsilon_t \prec 0] \circ \varepsilon_t$ ($I[\bullet]$ is a k*1 indicator function which takes on value 1 if the argument is true and 0 otherwise, while "$\circ$" indicates the Hadamard product) and $\bar{N} = E\left[n_t n'_t\right]$. Eq. (5) is the AG-DCC model. In order the Q_t to be positive definite for all possible realizations, the intercept, $\bar{P} - A'\bar{P}A - B'\bar{P}B - G'\bar{N}G$ must be positive semi-definite and the initial covariance matrix Q_0 positive definite.

3. Data

For our empirical analysis, we obtained a dataset of daily local currency denominated Morgan Stanley Capital International (MSCI) and J.P. Morgan emerging bond market index (ELMI) total returns from stock and bond markets respectively from 9 emerging countries.[3] The emerging markets examined are from various regions around the world- 2 from Latin America (Argentina and Mexico), 3 from the Middle East and Africa (Israel, South Africa, Turkey), 2 from Asia (Indonesia, Thailand), 2 from Emerging Europe (Czech Republic and Greece) and two developed countries from Asia (Hong Kong and Singapore). We also include two world indices: The MSCI world stock market index, which measures the equity market performance of developed markets and the J.P. Morgan global bond index (GBI Global), which includes local debt indices and tracks fixed rate issuances from high-income countries spanning North America, Europe, and Asia. For the analysis of the subprime crisis in EMEs, we expand the dataset by including the MSCI U.S. Broad Market and FTSE U.S. Government bond indexes as representatives for the U.S. equity and bond markets, respectively.

The sample period is from January 2, 1994 till end of December 2008. Our coverage in time and across countries was limited by the availability of all data required. We split our data as follows: (i) Asian crisis: 1997; (ii) Russian default: 1998; (iii) Argentine crisis: 1999–2000; (iv) Subprime crisis: July 2007–July 2008. We use Czech Republic as the crisis country for the Russian default, Thailand as the crisis country for the Asian crisis, Argentina as the crisis country for the Argentine turmoil and USA as the source of contagion of the subprime crisis.[4] The Johansen cointegration and VEC analysis is conducted during the entire period as well as each crisis period.

Then, using the AG-DCC model, we split the estimation procedure into subgroups in order to examine how similar the patterns of correlations (in the second moments) are between each crisis country and all other countries during stable and crises periods and compare the impact and the magnitude of spread of the crises in each individual emerging country's stock and bond market. One difficulty in testing for contagion during these periods is that there is no a single event to act as a definite catalyst behind the turmoil periods. Therefore, this study takes into account a stable period before and after the actual crash, using all historical information contained in pre and post-crash data (excluding the period 1/1/2001 till 27/09/2002 due to the dot-com bubble collapse).[5]

4. Empirical results

4.1. Johansen cointegration test and VECM results

Table 1 shows the cointegration results among bond and stock markets during the entire examined period (1994–2008).[6] We find there are three cointegrating vectors among bond markets (Panel A). Moreover, results reported in Panel B show that there is one cointegrating vector among the thirteen stock indices during the period 1994–2008.[7]

Table 2 shows the bilateral cointegration results between Czech stock (bond) index and the rest of the eleven stock (bond) indices, when Czech Republic is considered as the source of contagion for

[3] Using indices expressed in local currencies restricts their changes to movements in the prices and avoids distorting the cointegration analysis results with devaluations of the exchange rates that took place in the regions. However, our findings are basically the same when the data are converted in dollars.

[4] Instead of using Russia as source of the Russian default, we include Czech Republic as the geographically close Eastern European emerging country affected by the Russian crisis in order to examine the contagion effect from a different perspective than other similar studies (e.g., Dungey et al., 2007).

[5] Following Forbes and Rigobon (2002), a sensitivity analysis shows that period definition (tranquil and turbulent period) does not affect the central results.

[6] Summary statistics for stock and bond returns for the entire sample period exhibit quite common characteristics of financial time series (asymmetries, fat tails, non-normality), so not presented here due to space limitations. The Augmented Dickey–Fuller, the Phillips–Perron and the KPSS unit root tests results indicate that the null hypothesis of a unit root in the levels cannot be rejected for stock and bond series, while a unit root in the first differences is rejected at the 5% significance level. The Zivot–Andrews unit root test produce results that conform to the outcomes of the other unit root tests. Results are available up on request.

[7] When we estimate a VECM model for the thirteen bond indices (not presented here due to space limitations), we find the error correction term or the speed of adjustment coefficient is significant at 5% level only for Argentina, Hong Kong, Indonesia, Israel, Turkey and the global bond index. Moreover, bond indices of Indonesia, South Africa, Argentina and Thailand are significant to lag values of the global bond index. The VECM results for the thirteen stock indices (not presented here but available up on request) show significant short run dynamics among the MSCI world index and all other indices. The speed of adjustment coefficient is significant for all the indices, except Singapore, Turkey, Czech Republic and Argentina.

Table 1
Cointegration results among markets (period: 1994–2008).

Null hypothesis	Trace	5% Critical value	Max. eigenvalue	5% Critical value
Panel A: bond markets				
$r=0$	443.1256*	285.1425	129.8821*	70.53513
$r\leq 1$	310.5189*	239.2354	70.8965*	64.50472
$r\leq 2$	244.2277*	197.3709	63.5821*	58.43354
$r\leq 3$	176.1412*	159.5297	59.5541*	52.36261
Panel B: stock markets				
$r=0$	309.8521*	285.1425	77.4289*	70.53513
$r\leq 1$	230.1076	239.2354	52.3968	64.50472
$r\leq 2$	173.7198	197.3709	42.8921	58.43354
$r\leq 3$	126.7491	159.5297	30.4998	52.36261

Based on the AIC criteria, the appropriate lag length is two.
* Indicates significance at 5% level

the Russian crisis (period 1998). Both the trace and maximum eigenvalue tests for the stock markets support the existence of a single cointegration relation among Czech Republic and Argentina and among Czech Republic and Mexico. Applying a bivariate VECM for the cointegrated stock index series during the Russian crisis (results not presented here), we report a bi-directional causality from Czech Republic to Argentina and a uni-directional causality relation among the Czech and Mexican stock indices. The bilateral cointegration results for the bond markets show that there is no cointegration among Czech Republic and the rest of the eleven bond indices. Finally, results from Table 3 show there is no cointegration between Argentina and the rest of the indices during the Argentine crisis (1999–2000), for both equity and bond markets. This implies that there is no effect from Argentina to the rest of the countries, implying the isolated nature of this shock.

Table 2
Bilateral cointegration results among stock and bond markets during the Russian default (Czech Republic as source of contagion, period: 1998).

	Null hypothesis	Trace	Max. eigenvalue
CZE-ARG	$r=0$	16.34061* (7.747714)	14.69340* (7.696364)
	$r=1$	1.647210 (0.051350)	1.647210 (0.051350)
CZE-GRE	$r=0$	2.587637 (7.873609)	2.569279 (7.098508)
	$r=1$	0.018358 (0.775101)	0.018358 (0.775101)
CZE-HKG	$r=0$	12.33394 (8.296197)	8.195857 (8.269915)
	$r=1$	4.138087 (0.026282)	4.138087 (0.026282)
CZE-IDO	$r=0$	10.13331 (14.89029)	8.409523 (13.48481)
	$r=1$	1.723786 (1.405480)	1.723786 (1.405480)
CZE-ISR	$r=0$	5.065372 (4.412772)	3.982600 (3.908101)
	$r=1$	1.082772 (0.504670)	1.082772 (0.504670)
CZE-MEX	$r=0$	17.45400* (7.517312)	14.30283* (5.226924)
	$r=1$	3.151175 (2.290389)	3.151175 (2.290389)
CZE-SAF	$r=0$	7.072160 (6.174850)	6.269249 (3.960253)
	$r=1$	0.802910 (2.214597)	0.802910 (2.214597)
CZE-SNG	$r=0$	11.66445 (5.202390)	7.004195 (4.811571)
	$r=1$	4.660253 (0.390819)	4.660253 (0.390819)
CZE-THA	$r=0$	10.14208 (6.461919)	6.321128 (6.047594)
	$r=1$	3.820954 (0.414325)	3.820954 (0.414325)
CZE-TUR	$r=0$	11.51544 (8.028608)	10.88011 (7.415837)
	$r=1$	0.635323 (0.612770)	0.635323 (0.612770)
CZE-MSCI (CZE-GBI)	$r=0$	5.930228 (7.583019)	4.724982 (5.558318)
	$r=1$	1.205246 (2.024701)	1.205246 (2.024701)

Each cell in columns 3 and 4 contains estimates for stock markets. In parentheses, results for bond markets. The 5% critical values for the trace test ($r=0$ and $r\leq 1$) are 15.49 and 3.84, respectively. The corresponding critical values for maximal eigenvalue test ($r=0$ and $r=1$) are 14.26 and 3.84, respectively.
* Indicates significance at 5% level.

Table 3
Bilateral cointegration results among stock and bond markets during the Argentine crisis (Argentina as source of contagion, period: 1999–2000).

	Null hypothesis	Trace	Max. eigenvalue
ARG-CZE	$r=0$	9.328611 (9.924522)	6.194739 (9.459278)
	$r=1$	3.133872 (0.465244)	3.133872 (0.465244)
ARG-GRE	$r=0$	8.097726 (5.647671)	6.163266 (4.732090)
	$r=1$	1.934460 (0.915580)	1.934460 (0.915580)
ARG-HKG	$r=0$	8.337589 (15.65539)	6.992472 (14.39948)
	$r=1$	1.345117 (2.255909)	1.345117 (2.255909)
ARG-IDO	$r=0$	6.800213 (5.654416)	5.239374 (5.011301)
	$r=1$	1.560839 (0.643116)	1.560839 (0.643116)
ARG-ISR	$r=0$	7.129805 (9.184582)	5.753647 (8.117269)
	$r=1$	1.376158 (1.067313)	1.376158 (1.067313)
ARG-MEX	$r=0$	6.483587 (8.274293)	4.743895 (7.305823)
	$r=1$	1.739692 (0.968470)	1.739692 (0.968470)
ARG-SAF	$r=0$	9.700488 (5.007629)	8.118164 (3.336004)
	$r=1$	1.582323 (1.671625)	1.582323 (1.671625)
ARG-SNG	$r=0$	8.238484 (7.379091)	5.444666 (6.380767)
	$r=1$	2.793818 (0.998324)	2.793818 (0.998324)
ARG-THA	$r=0$	9.392167 (5.084109)	8.411564 (3.599043)
	$r=1$	0.980603 (1.485066)	0.980603 (1.485066)
ARG-TUR	$r=0$	4.248096 (8.394001)	2.842473 (7.531359)
	$r=1$	1.405623 (0.862642)	1.405623 (0.862642)
ARG-MSCI (ARG-GBI)	$r=0$	10.75535 (4.529434)	9.952891 (3.474453)
	$r=1$	0.802456 (1.054981)	0.802456 (1.054981)

Each cell contains estimates for stock markets. In parentheses, results for bond markets. The 5% critical values for the trace test ($r=0$ and $r\leq 1$) are 15.49 and 3.84, respectively. The corresponding critical values for maximal eigenvalue test ($r=0$ and $r\leq 1$) are 14.26 and 3.84, respectively.

Table 4
Bilateral cointegration results among stock and bond markets during the Asian crisis (Thailand as source of contagion, period: 1997).

	Null hypothesis	Trace	Max. eigenvalue
THA-ARG	$r=0$	15.46552 (5.794135)	9.961485 (5.742427)
	$r=1$	5.504035 (0.051708)	5.504035 (0.051708)
THA-CZE	$r=0$	16.91988* (8.169759)	12.82416 (6.874044)
	$r=1$	4.095723 (1.295716)	4.095723* (1.295716)
THA-GRE	$r=0$	6.732360 (2.881871)	6.536086 (2.604408)
	$r=1$	0.196274 (0.277464)	0.196274 (0.277464)
THA-HKG	$r=0$	11.18520 (4.181295)	7.046959 (3.903292)
	$r=1$	4.138245 (0.278003)	4.138245 (0.278003)
THA-IDO	$r=0$	13.37650 (17.96535)	9.181826 (12.93383)
	$r=1$	4.194677 (5.031518)	4.194677 (5.031518)
THA-ISR	$r=0$	9.352183 (9.721034)	6.761338 (8.320861)
	$r=1$	2.590846 (1.400172)	2.590846 (1.400172)
THA-MEX	$r=0$	16.04609* (7.213187)	12.79903 (4.143231)
	$r=1$	3.247065 (3.069956)	3.247065 (3.069956)
THA-SAF	$r=0$	15.61665* (9.034477)	10.43744 (8.237923)
	$r=1$	5.179212 (3.343849)	5.179212 (3.343849)
THA-SNG	$r=0$	12.47241 (9.527945)	7.811649 (8.237923)
	$r=1$	4.660757 (1.290022)	4.660757 (1.290022)
THA-TUR	$r=0$	16.14843* (14.64723)	11.93353 (14.17667)
	$r=1$	4.214900 (0.470555)	4.214900 (0.470555)
THA-MSCI(THA-GBI)	$r=0$	7.550899 (7.553979)	6.228114 (4.472924)
	$r=1$	1.322785 (3.081055)	1.322785 (3.081055)

Each cell in columns 3 and 4 contains estimates for stock markets. In parentheses, results for bond markets. The 5% critical values for the trace test ($r=0$ and $r\leq 1$) are 15.49 and 3.84, respectively. The corresponding critical values for maximal eigenvalue test ($r=0$ and $r=1$) are 14.26 and 3.84, respectively.

* Indicates significance at 5% level.

Table 5
Number of cointegrating vectors among stock and bond markets during the subprime crisis (USA as source of contagion, period: July 2007–July 2008).

	Cointegrating vectors	
	Stock markets	Bond markets
USA-ARG	$r=1$	$r=0$
USA-CZE	$r=1$	$r=1$
USA-GRE	$r=1$	$r=1$
USA-HKG	$r=1$	$r=1$
USA-IDO	$r=1$	$r=1$
USA-ISR	$r=1$	$r=1$
USA-THA	$r=1$	$r=1$
USA-MEX	$r=1$	$r=0$
USA-SAF	$r=1$	$r=1$
USA-SNG	$r=1$	$r=1$
USA-TUR	$r=1$	$r=1$
USA-MSCI(USA-GBI)	$r=3$	$r=1$

The number of cointegrating vectors for both stock and bond markets are obtained using the Johansen procedure (trace and maximal eigenvalue tests). Tests statistics are not presented but are available up on request.

Table 4 reports the bilateral cointegration results among Thailand stock (bond) market and the rest of the stock (bond) markets, considering Thailand as the source of the Asian crisis (period 1997). We reject the null hypothesis (no cointegration) among the stock markets of Thailand and Czech Republic, Thailand and Mexico, Thailand and South Africa, Thailand and Turkey. According to bivariate VECM results for the cointegrated stock index series (not reported here), we find significant error correction terms in the case of Thailand and Turkey, one error correction term in the case of Thailand and Czech Republic and a uni-directional causality among Thailand and Mexico. The results for bond markets show that there is no cointegration between the bond market of Thailand and the rest of the eleven bond indices. Given that there are no linear combinations of bond returns that are stationary, there is no error correction representation. This implies that investors can gain substantial long-term diversification benefits.

Table 5 reports the number of cointegrating vectors among USA (stock and bond markets) and the rest of stock (bond) markets using the Johansen procedure (MSCI U.S. Broad Market and FTSE U.S. Government bond indexes considered as the source of the subprime crisis for each asset market). We find one cointegrating vector between USA and all EMEs for both stock and bond markets, with the exception of the MSCI world equity index ($r=3$), Argentine and Mexican bond markets (no cointegrating relationship). Bivariate VECM results (not reported here), show significant error correction terms in almost all pairs, a bi-directional causality between U.S. and almost all EMEs, and a uni-directional causality among U.S. and the two world indices.

4.2. *Estimates of the AG-DCC model*

4.2.1. *Results for emerging market crises*

Table 6 shows non-parametrically the presence of asymmetries in conditional second moments for stock and bond markets during the Russian crisis. Results indicate average conditional correlations between Czech Republic and other countries for both stock and bond markets are higher during the negative shock than stable periods, supporting the contagion phenomenon. The effects of contagion for both stock and bond markets are strongest on the Latin American (Argentina and Mexico) and the Asian emerging and developed countries (Thailand, Indonesia, Hong Kong and Singapore). Moreover, the contagion effects in equities differ to those reported in bond markets for this period. Specifically, the spread of the crisis is larger in magnitude among stock rather than bond markets. This suggests that the nature of particular assets or asset markets may hold important information on the transmission of shocks. Finally, the Russian default seems to have a global impact on both stock and bond markets, since the average correlation between the Czech Republic and the two world indices is significantly

Table 6
Estimates of conditional average correlations for stock and bond markets during the Russian default (Czech Republic as source of contagion vs. others).

		Crisis period (1998)	Stable period
Stock markets			
CZE	ARG	0.50*	0.43*
	GRE	0.35*	0.30*
	HKG	0.53*	0.33*
	IDO	0.45*	0.28
	ISR	0.34*	0.33*
	MEX	0.51*	0.43*
	SAF	0.33*	0.20
	SNG	0.39*	0.23*
	THA	0.36*	0.22*
	TUR	0.32*	0.29
	MSCI	0.48*	0.32*
Bond markets			
CZE	ARG	0.44*	0.35
	GRE	0.30*	0.26*
	HKG	0.48*	0.36*
	IDO	0.40*	0.25
	ISR	0.30*	0.27*
	MEX	0.41*	0.36*
	SAF	0.24*	0.21*
	SNG	0.31*	0.24*
	THA	0.32*	0.20
	TUR	0.29*	0.26*
	GBI	0.32*	0.23*

This table reports conditional correlations during crisis and stable periods. Estimates are obtained using the AG-DCC model. We estimate the conditional average correlations between Czech Republic- the crisis country (equity and bond market) and all other countries. *T*-test statistics are of tests of equivalent correlation with the crisis country (Czech Republic) equity index and bond return between the turmoil and stable periods.

* The rejection of the null hypothesis against the one-sided alternative that during the turmoil period correlation increases and contagion spreads is tested at the 5% significance levels.

higher during the crisis than stable periods. Our findings are in line with earlier work of Dungey et al. (2007).

Table 7 reports conditional average correlations between Argentina and each one of the other countries for both stock and bond markets. The Argentine crisis exhibit strong isolation characteristics, since there is not any significant increase in correlation from stable to crisis period among Argentina and all other countries for both stock and bond markets. This is in line with the absence of any long-run cointegrating relationship among the examined markets during this specific shock reported in this paper and the work of Boschi (2005), which shows the lack of spillovers from the Argentine crisis applying conventional approaches.

While both the Russian and Argentinean defaults were caused by weak fiscal policy and governments' inability to reduce public and external debt, their differences in crisis dynamics could be explained by two factors. First, changes in investors' behavior during the 1990s and early 2000. Non-dedicated investors (high-yield funds, hedge funds, investment banks), who dominated investments in emerging financial markets during the 1990s, have become much less significant players, leaving emerging markets after banks' declining willingness to extend credit to them in the aftermath of the Russian default and the crash of Long-Term Capital management (LTCM) in 1998 (IMF International Capital Markets, August 2001). Second, the nature of each crisis seems to matters. While the Argentine crisis was largely anticipated by the markets, Russian crisis had a big element of surprise. In investors' minds Russia was too important to collapse. However, investors were more properly equipped to navigate through the Argentine turmoil, as they learned their lesson from the Russian crisis.

Table 8 reports conditional average correlations between Thailand and each one of the other countries for both stock and bond markets. All estimated correlations are higher during the Asian turmoil than stable periods, with the exception of the world stock and bond indices. Moreover, the shocks

Table 7
Estimates of conditional average correlations for stock and bond markets during the Argentine crisis (Argentina as source of contagion vs. others).

		Crisis period (1999–2000)	Stable
Stock markets			
ARG	CZE	0.20	0.24
	GRE	0.12	0.23*
	HKG	0.23*	0.33*
	IDO	0.25	0.28
	ISR	0.34	0.33*
	MEX	0.51*	0.56*
	SAF	0.13*	0.20
	SNG	0.19*	0.23
	THA	0.16*	0.22
	TUR	0.19	0.19*
	MSCI	0.22*	0.32*
Bond markets			
ARG	CZE	0.14*	0.23
	GRE	0.56*	0.45*
	HKG	0.25	0.26*
	IDO	0.23	0.25
	ISR	0.22*	0.32*
	MEX	0.41*	0.46*
	SAF	0.18*	0.22
	SNG	0.20*	0.27*
	THA	0.19	0.20
	TUR	0.18*	0.22*
	GBI	0.32*	0.35*

This table reports conditional correlations during crisis and stable periods. Estimates are obtained using the AG-DCC model. We estimate the conditional average correlations between Argentina- the crisis country (equity and bond market) and all other countries. *T*-test statistics are of tests of equivalent correlation with the crisis country (Argentina) equity index and bond return between the turmoil and stable periods.

* The rejection of the null hypothesis against the one-sided alternative that during the turmoil period correlation increases and contagion spreads is tested at the 5% significance levels.

originated from Thailand are transmitted with a larger magnitude through the Thai equity market than the Thai bond market. However, the magnitude of the Asian crisis spread is higher between the Asian countries rather than the other emerging countries of different regions. This evidence does not agree with the “no contagion” conclusion of Forbes and Rigobon (2002), but is in line with the findings of other studies on Asian crisis (e.g., Sheng and Tu, 2000; Chiang et al., 2007).

Trade and bank linkages together with financial vulnerabilities seem to account for the different contagious effects of the Asian crisis compared to the Argentine default (IMF, Global Financial Stability Report, 2000; BIS, Financial Stability Review, December 2001; Bank of England, Financial Stability Review, June 2001 and 2002). Asian economies, which experienced the most severe spillovers form the Thai crisis, had both relatively strong trade links with Thailand (similar patterns in export destinations prior to the crisis) and large current account deficits. Many also had managed exchange rate systems and had seen appreciations in their real exchange rate positions prior to the crisis (Corsetti et al., 1999). They also tended to have strong banking sector dependencies on Japan, which may have interacted with generally low reserve coverage of short-term debt. East Asian countries enjoyed substantial short term capital inflows amounted to five to ten percent of gross domestic product during the 1990s. However, the event of the crisis increased rapidly the speed at which these funds exit the Asian economies (according to the BIS, capital outflows from Asia amounted to $102 billion in the second half of 1997).

Those linkages between Argentina and EMEs were much lower, while many EMEs have adopted floating exchange rate regimes that may provide an additional shield to spillovers. The ruble’s massive devaluation followed by sovereign debt default boosted emerging market risk and suppressed commodity exports from emerging markets to Russia. However, the default episode and the collapse

Table 8
Estimates of conditional average correlations for stock and bond markets during the Asian crisis (Thailand as source of contagion vs. others).

		Crisis period (1997)	Stable period
Stock markets			
THA	ARG	0.42*	0.33
	CZE	0.30*	0.27
	GRE	0.39*	0.30
	IDO	0.70*	0.54*
	ISR	0.41*	0.33
	MEX	0.39*	0.31
	SAF	0.40*	0.30
	SNG	0.62*	0.45*
	HKG	0.72*	0.53*
	TUR	0.46*	0.30*
	MSCI	0.26*	0.31*
Bond markets			
THA	ARG	0.34*	0.29
	CZE	0.27*	0.23
	GRE	0.32*	0.23*
	IDO	0.65*	0.44*
	ISR	0.34*	0.25
	MEX	0.28*	0.22
	SAF	0.35*	0.27
	SNG	0.56*	0.45*
	HKG	0.70*	0.53*
	TUR	0.33*	0.28
	GBI	0.15*	0.18*

This table reports conditional correlations during crisis and stable periods. Estimates are obtained using the AG-DCC model. We estimate the conditional average correlations between Thailand- the crisis country (equity and bond market) and all other countries. *T*-test statistics are of tests of equivalent correlation with the crisis country (Thailand) equity index and bond return between the turmoil and stable periods.

* The rejection of the null hypothesis against the one-sided alternative that during the turmoil period correlation increases and contagion spreads is tested at the 5% significance levels.

of the U.S. hedge fund LTCM within weeks of its onset led to a reappraisal of credit and sovereign risks across financial markets in both developed and emerging countries, high interest rates both in the U.S. and other economies, and widespread fears of a liquidity crash.

4.2.2. Results for the subprime crisis

Table 9 reports conditional average correlations between USA and each one of the other countries for both stock and bond markets during stable and crisis periods. Results show that the subprime crisis is spread through both stock and bond markets in all regions. Asian developed and emerging countries were most hit, while the subprime crisis has a global impact, since average correlations between U.S. stock and bond indices and the two world indices are significantly increased from stable to crisis period. Moreover, the propagation of the crisis is much stronger through EMEs stock markets than bond markets. In the wake of the crisis, a slower world economy, net capital outflows, bank vulnerabilities due to short-term international financing exposure and risky lending practices and insufficient international reserves or the absence of fiscal surpluses may account for the contagious effects in EMEs. Our results are in line with the rejection of the decoupling hypothesis supported by Dooley and Hutchison (2009) and Aloui et al. (2011).

However, emerging countries in Latin America (Argentina and Mexico) and Africa and Middle East (Israel, South Africa and Turkey) seem that were least affected by the subprime crisis, since their pairwise correlations with US indices (for both stock and bond markets) increased slightly from stable to crisis period. The lower vulnerability of Latin American countries to the subprime crisis implies that this continent was the most prepared, given its history of financial crises. Latin America countries have built up to 400 billion dollars in international reserves and they have substantially reduced their

Table 9
Estimates of conditional average correlations for stock and bond markets during the subprime crisis (USA as source of contagion vs. others).

		Crisis period (2007–2008)	Stable
Stock markets			
U.S.	ARG	0.64*	0.62*
	CZE	0.72*	0.57
	GRE	0.82	0.68*
	HKG	0.83*	0.73*
	IDO	0.77*	0.39*
	ISR	0.55	0.53*
	MEX	0.61*	0.59*
	SAF	0.38*	0.37*
	SNG	0.89*	0.74*
	THA	0.71*	0.42*
	TUR	0.40*	0.39*
	MSCI	0.88*	0.72*
Bond markets			
U.S.	ARG	0.40*	0.39*
	CZE	0.44*	0.36
	GRE	0.58*	0.47*
	HKG	0.55	0.46*
	IDO	0.32	0.26
	ISR	0.44*	0.42*
	MEX	0.33*	0.31*
	SAF	0.27*	0.25*
	SNG	0.39*	0.33*
	THA	0.35*	0.29
	TUR	0.26*	0.25*
	GBI	0.57*	0.48*

This table reports conditional correlations during crisis and stable periods. Estimates are obtained using the AG-DCC model. We estimate the conditional average correlations between USA- the crisis country (equity and bond market) and all other countries. *T*-test statistics are of tests of equivalent correlation with the crisis country (USA) equity index and bond return between the turmoil and stable periods.

* The rejection of the null hypothesis against the one-sided alternative that during the turmoil period correlation increases and contagion spreads is tested at the 5% significance levels.

dollar-denominated debt, especially within the banking system. In the wake of the financial crisis, they swiftly depreciated their currencies, while many of these saved a considerable amount of their tax income on extra revenue from commodity bonanza at the turn of the century. In the same line, the strong macroeconomic and financial conditions in the three EMEs in Africa and Middle East during the past years seem to reduce their vulnerability to the recent financial meltdown.

4.2.3. Robustness tests

Table 10 reports the results from four different parameterizations of DCC models estimated in order to confirm whether we should adopt a symmetric or an asymmetric model (during the three contagious episodes). All parameters are significantly different from zero. The g_i^2 term in the asymmetric model is always higher than zero, implying the presence of asymmetric movements. Furthermore, the asymmetric b_i^2 term is always higher than the symmetric b_i^2, providing further support to the use of the asymmetric DCC model in this study. Therefore, the results show that conditional correlations among the crisis country (stock and bond markets) and all others are much greater during extreme downside moves than upside moves. The estimated unconditional correlations among markets, not reported here for brevity, show that they increase significantly during the three contagious crises. Moreover, conditional correlations are substantially greater than unconditional correlations, supporting the presence of asymmetric contagion during the Asian, Russian and subprime crises.

Table 10
DCC GARCH models.

Crisis country vs. all others	Symmetric model		Asymmetric model		
	a_i^2	b_i^2	a_i^2	g_i^2	b_i^2
Stock markets:					
Czech	0.0018*	0.9634*	0.0015*	0.0011*	0.9695*
Thailand	0.0016*	0.9601*	0.0012*	0.0009*	0.9674*
U.S.	0.0171*	0.9701*	0.0162*	0.0024*	0.9789*
Bond markets:					
Czech	0.0018*	0.9634*	0.0015*	0.0011*	0.9695*
Thailand	0.0016*	0.9601*	0.0012*	0.0009*	0.9674*
U.S.	0.0112*	0.9532*	0.0131*	0.0013*	0.9746*

Four different parameterizations are estimated for the dynamics of correlation during the three crisis periods (Russian default in 1998, Asian crisis in 1997 and the subprime crisis of 2007–2008):
(a) A standard scalar DCC model with no asymmetric terms of the following form: $Q_t = (1 - a - b)\bar{P} + a\varepsilon_{t-1}\varepsilon'_{t-1} + bQ_{t-1}$.
$P_t = Q_t^{*-1} Q_t Q_t^{*-1}$.
where, $\bar{P} = E[\varepsilon_t \varepsilon'_t]$ and α and b are scalars such that $\alpha + b < 1$. $Q_t^* = \left[q_{iit}^*\right] = [\sqrt{q_{iit}}]$ is a diagonal matrix with the square root of the ith diagonal element of Q_t on its ith diagonal position. As long as Q_t is positive definite, Q_t^* is a matrix which guarantees $P_t = Q_t^{*-1} Q_t Q_t^{*-1}$ is a correlation matrix with ones on the diagonal and every other element ≤ 1 in absolute value.
(b) A symmetric diagonal version (where matrices A and B are diagonal and matrix G is set equal to zero) of the following form: $Q_t = (\bar{P} - A'\bar{P}A - B'\bar{P}B - G'\bar{N}G) + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + G'n_{t-1}n'_{t-1}G + B'Q_{t-1}B$.
(c) An asymmetric scalar DCC model of the following form: $Q_t = (\bar{P} - a^2\bar{P} - b^2\bar{P} - g^2\bar{N}) + a^2\varepsilon_{t-1}\varepsilon'_{t-1} + g^2 n_{t-1}n'_{t-1} + b^2 Q_{t-1}$.
A sufficient condition for Q_t to be positive definite is that the matrix in parentheses is positive semi-definite. A necessary and sufficient condition for this to hold is $\alpha^2 + b^2 + \delta g^2 < 1$, where δ is the maximum eigenvalue $[\bar{P}^{-1/2}\bar{N}\bar{P}^{-1/2}]$.
(d) The full diagonal version of a scalar A-DCC (the matrices A, B and G are assumed to be diagonal) of the following form: $Q_t = \bar{P} \circ (ii' - aa' - bb') - \bar{N} \circ gg' + aa' \circ \varepsilon_{t-1}\varepsilon'_{t-1} + gg' \circ n_{t-1}n'_{t-1} + bb' \circ Q_{t-1}$.
where, i is a vector of ones and a, b and g are vectors containing the diagonal elements of matrices A, B and G, respectively.
* Indicates significance at the 5% level.

5. Concluding remarks

This paper investigates the correlated-information channel as a contagion mechanism for three emerging market crises of the late 1990s, as well as the subprime crisis of 2007, using data from equity and bond markets of EMEs from various regions around the world, USA and 2 global stock and bond indices. Through conventional cointergation and VEC analysis, we report long and short run dynamics only for stock markets during the Russian and Asian crises, for both stock and bond markets during the subprime crisis, while the Argentine crisis has no impact on any of the examined financial markets.

To provide a more robust analysis of financial contagion, we also examine conditional correlation dynamics into a time varying asymmetric framework, applying the recently developed AG-DCC model. This empirical analysis elucidates how vulnerable emerging financial markets are to both emerging and global shocks and displays differences in crises dynamics that could be attributed to several factors. Results confirm the existence of asymmetric contagion to EMEs and globally only for the Russian default and the subprime crisis. Both crises hit those economies regardless of their economic integration, since cross-market correlation dynamics are driven by behavioral reasons, due to shifting investor sentiment (increased risk aversion), causing significant changes in the emerging countries' financial structures. The other government default episode (Argentine crisis) examined in this paper has an isolated nature, due to the lower macroeconomic and financial vulnerability to shocks of emerging economies with close trade and financial ties to Argentina and the shifts in the emerging economies investor base due to earlier emerging market crises. On the other hand, the asymmetric contagious effects of the Asian crisis display strong intra-regional characteristics, due to the growing share of investment and trade in the region and the more common monetary policy followed by the Asian countries after the October crash of 1987.

Our findings have important implications for international investors and portfolio managers. Evidence on contagion implies that diversification sought by investing in multiple markets from different regional blocks is likely to be lower when it is most desirable. As a result, an investment strategy focused solely on international diversification seems not to work in practice during turmoil periods. On the other hand, evidence provided on the lower spread of all four crises through bond markets indicate that, in times of distress, any potential benefits from international diversification are greater for the bond investors than the stock investors. Finally, since countries and financial markets react differently to sovereign shocks, combining bonds and stocks from different emerging economies could provide advantages over debt-only or equity-only portfolios.

The results also provide useful implications regarding the ability of policy makers and multi-lateral organizations to insulate or at least attenuate an economy from contagious effects. The intra-regional contagious effects of the Asian crisis indicate the effectiveness of the co-ordinated rescue packages from the International Monetary Fund (IMF) to the most affected Southeast and East Asian (Thailand, Indonesia and South Korea). The rescue packages were of a considerable sum, as of July 1998 the combined value of the IMF rescue packages amounted to a total of $100 billion. Regarding the Argentine crisis, policy initiatives by both the IMF and EMEs (e.g., increased EME data dissemination, floating rate regimes) following previous crises have led to improvements in country surveillance. In the case of the Russian crisis, the high interest rates both in the U.S. and other economies after the shock led to widespread fears of a liquidity crash and raised questions about policy performance and coordination, despite the bailout of LTCM organized by the U.S. Fed and the followed expansionary monetary policy on a global base.

Finally, the subprime crisis raised the need for a revamped international financial architecture. The global contagious effects of this crisis and the rejection of the decoupling hypothesis for the EMEs question the resilience and sustainability of emerging-market policy performance. It seems that strong economic indicators in many EMEs before the crisis (high growth rates, massive foreign exchange reserves, balanced budgets) were not enough to decouple them from the crisis, because of their cyclicality and endogeneity. A consequence of the contagion on EMEs would be the redirection of development loans by the World Bank, the IMF, and the regional development banks to the public sector, since those funds had been crowded out by private-sector lending throughout the boom decade (World Bank, Global Development Finance, 2008).

References

Acharya, V., Pedersen, L., 2005. Asset pricing with liquidity risk. Journal of Financial Economics 77, 375–410.

Allen, F., Gale, D., 2000. Financial contagion. Journal of Political Economy 108, 1–33.

Aloui, R., Ben Aossa, M.S., Nguyen, D.K., 2011. Global financial crisis, extreme interdependences, and contagion effects: the role of economic structure? Journal of Banking & Finance 35, 130–141.

Baig, T., Goldfajn, I., 1999. Financial market contagion in the Asian crisis. IMF Staff Papers 46.

Baur, D., 2003. Testing for contagion—mean and volatility contagion. Journal of Multinational Financial Management 13, 405–422.

Bekaert, G., Harvey, C.R., Ng, A., 2005. Market integration and contagion. Journal of Business 78, 39–68.

Bollerslev, T., 1986. Generalized autoregressive conditional heteroskedasticity. Journal of Econometrics 31 (3), 307–327.

Boschi, M., 2005. International financial contagion: evidence from the Argentine crisis of 2001–2002. Applied Financial Economics 15, 153–163.

Calvo, S., Reinhart, C.M., 1996. Capital flows to Latin America: is there evidence of contagion effects? In: Calvo, G., Goldstein, M., Hochreiter, E. (Eds.), Private Capital Flows to Emerging Markets. Institute for International Economics, Washington, DC.

Cappiello, L., Engle, R.H., Sheppard, K., 2006. Asymmetric dynamics in the correlations of global equity and bond returns. Journal of Financial Econometrics 4, 537–572.

Chiang, T.C., Jeon, B.N., Li, H., 2007. Dynamic correlation analysis of financial contagion: evidence from the Asian Markets. Journal of International Money and Finance 26, 1206–1228.

Corsetti, G., Pesenti, P., Roubini, N., 1999. What caused the Asian currency and financial crisis? Japan and the World Economy 11, 305–373.

Dooley, M., Hutchison, M., 2009. Transmission of the U.S. subprime crisis to emerging markets: evidence on the decoupling–recoupling hypothesis. Journal of International Money and Finance 28, 1331–1349.

Dornbusch, R., Park, Y., Claessens, S., 2000. Contagion: understanding how it spreads. The World Bank Research Observer 15, 177–197.

Dungey, M., Fry, R., Gonzalez-Hermosillo, B., Martin, V.L., 2007. Contagion in global equity markets in 1998. The effects of the Russian and LTCM crises. The North American Journal of Economics and Finance 18, 155–174.

Engle, R.F., 2002. Dynamic conditional correlation - a simple class of multivariate GARCH models. Journal of Business and Economic Statistics 20, 339–350.

Forbes, K., 2000. The Asian flu and Russian virus: firm-level evidence on how crises are transmitted internationally. Working paper, National Bureau of Economic Research.

Forbes, K.J., Rigobon, R., 2002. No contagion, only interdependence: measuring stock market comovements. Journal of Finance 57, 2223–2261.

Gelos, G., Sahay, R., 2000. Financial Market Spillover in Transition Economies. IMF Working Paper 00/71.
Glick, R., Rose, A., 1999. Contagion and trade: why are currency crises regional? Journal of International Money and Finance 18, 603–617.
Johansen, S., 1991. Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models. Econometrica 59, 1551–1580.
Johansen, S., Juselius, K., 1990. Maximum likelihood estimation and inference on cointegration with application to the demand for money. Oxford Bulletin of Economics and Statistics 52, 169–210.
Kaminsky, G., Reinhardt, C., Vegh, C., 2003. The unholy trinity of financial contagion. Journal of Economic Perspectives 17, 51–74.
Kenourgios, D., Samitas, A., Paltalidis, N., 2011. Financial crises and stock market contagion in a multivariate time-varying asymmetric framework. Journal of International Financial Markets, Institutions & Money 21 (1), 92–106.
King, M., Wadhwani, S., 1990. Transmission of volatility between stock markets. Review of Financial Studies 3, 5–33.
Okimoto, T., 2008. New evidence of asymmetric dependence structures in international equity markets. Journal of Financial and Quantitative Analysis 43, 787–815.
Pelletier, D., 2006. Regime switching for dynamic correlation. Journal of Econometrics 131, 445–473.
Rodriquez, J.C., 2007. Measuring financial contagion: a copula approach. Journal of Empirical Finance 14, 401–423.
Sheng, H.-C., Tu, A.H., 2000. A study of cointegration and variance decomposition among national equity indices before and during the period of the Asian financial crisis. Journal of Multinational Financial Management 10, 345–365.