The Effects of Contagion during the Global Financial Crisis in Government-Regulated and Sponsored Assets in Emerging Markets

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by

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in

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Abstract

The effects of financial contagion during the Global Financial Crisis (GFC) have been extensively studied in the finance literature. One of the key issues is the devastating effect of the crisis on wealth and asset prices. However, an important difference between this crisis and past crises was the relatively small and short duration of the effects of the crisis on emerging markets. Of particular interest was the resilience of government-regulated and -sponsored assets such as pension funds, state-owned enterprises, and international and local-currency government bonds. This thesis contributes to the literature on the effects of financial contagion by analysing four cases of government-regulated and -sponsored assets during different episodes of the GFC. The second chapter analyses contagion from US equity markets to emerging-market autarchic assets (Colombian private pension funds) during different episodes of the GFC. In this paper we test for contagion via changes in correlation between financial asset returns and via additional volatility spillovers. We propose a DCC-GJR GARCH framework where the S&P 500 is the source of transmission of contagion to the autarchic asset. We find no evidence of contagion measured as significant changes in correlation during the first two phases of the crises. In Chapter 3 we extend our analysis to government-sponsored assets (state-owned enterprises (SOEs)) and argue that these assets account for a substantial and increasing fraction of global foreign direct investment. While emerging-economy SOEs are often vehicles for state-directed economic growth policy, the performance of SOEs compared with private enterprises is an open question, particularly during crisis periods. We estimate a four-factor model of SOE returns of the BRIC economies for the period 2000–12 and show that certain SOEs offered some protection to investors during the financial crisis of 2007–09. We use quantile regressions since this approach is robust to the presence of outliers and their impact on the

factors during crisis periods. The results obtained in this chapter provide empirical evidence for the special role of the state in protecting and stabilising state-owned enterprises. In Chapter 4 we analyse the effects of the GFC on government bonds. For this objective we use propensity matching estimation to measure the effect of the GFC on sovereign spreads using data from 43 countries. We estimate general underlying factor models allowing for multiple channels of contagion transmission then use estimates to select matching non-crisis benchmarks for nine portfolios of sovereign bonds. We found no significant changes in spreads on portfolios of local-currency emerging-market debt during the GFC. Finally, in Chapter 5 we use high-frequency Colombian government bond data and perform an event study on high-frequency data to measure the effect of the news originating from the GFC via a market-transmission mechanism. In order to avoid confounding effects, we compare the impact of news originating from the GFC with global, regional, and local news. Our results make an interesting contribution to understanding the extent of the resilience of emerging markets under the postulates of the coupling/decoupling hypothesis and market integration.

Keywords: emerging markets, global financial crisis, regulation, stated-owned enterprises, pension funds, government bonds.

JEL classification: C5, G1, G2, G14, G15, G28, G38

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Chapter 1 Introduction

The effects of financial contagion during the Global Financial Crisis (GFC) have been extensively studied in the finance literature. One of the key issues is the devastating effect of the crisis on wealth and asset prices. However, an important difference between this crisis and past crises was the relatively small and short duration of the effects of the crisis on emerging markets. Dooley and Hutchison (2009) were the first to find evidence in support of the decoupling hypothesis of emerging markets during the early phases of the crisis. Since then the hypothesis has been tested by other researchers (for recent surveys see: Beirne and Gieck, 2014; Köksal and Orhan, 2013).

In this thesis we contribute to this body of knowledge by focusing on the effects of the crisis on government-regulated and -sponsored assets such as pension funds, state-owned enterprises, and international and local-currency government bonds. By doing this we hope to clarify the role that governments play in preventing loss of wealth by insulating their domestic markets, either indirectly by regulation, as in the case of Colombian pension funds (Chapter 2), or by direct ownership, as in the case of state-owned enterprises (Chapter 3). In both chapters we hypothesise that regulation and ownership are the main drivers behind decoupling and recoupling during different episodes of the GFC. In Chapter 4 we further refine the issue of common transmission channels for financial contagion during the GFC at the global, regional, and local levels and analyse the evidence of decoupling/recoupling in different phases of the GFC on sovereign bonds. Finally, in Chapter 5 we extend our analysis using high-frequency data on local Colombian currency bonds to test the impact and the effect of GFC events on prices and further corroborate the previous chapter's findings at a level of detail rarely seen using data from emerging markets. In the following subsections we provide more detail to the general content and method employed in each chapter.

1.1. Can financial autarchy prevent contagion? The case of Colombian pension funds during the subprime, financial, and sovereign debt crises.

The benefits and costs of regulation have been intensely debated in the context of financial crises. A key issue is the extent to which regulators and governments should attempt to quarantine domestic assets from external shocks. Chapter 2 analyses contagion originating from US equity markets during the Global Financial Crisis (GFC) as it impacted on regulated assets in emerging-economy pension funds. In this chapter we divide the GFC into three contiguous episodes: the subprime, credit crunch (CCC) and European sovereign debt (ESD) crises. We define financial autarchy here as an investment where the local investors have a restricted choice in portfolio selection, and investment in the financial autarchic asset is legally mandatory. It is within this definition that Colombian pension funds (the autarchic asset) provide an interesting case for testing for evidence of contagion.

In order to test for contagion we propose a DCC-GJR GARCH framework based on Glosten et al. (1993) where the S&P 500 is the source of transmission of contagion to the autarchic asset. Finally, we also propose the use of quantile regressions as a way to test the magnitude of the contagion effects derived from a time-varying systematic risk measure during different crisis episodes.

1.2. The performance of state-owned enterprises in BRIC countries during the GFC

In recent years there has been a fundamental change in foreign direct investment (FDI) in emerging markets as these countries shift from being recipients to becoming investors in their own right. UNCTAD (2012) estimates that for the first time in 2010, while developing countries received more than 50 per cent of total global FDI, 30 per cent of global FDI is coming directly from emerging markets. It is in this context that emerging-market state-owned enterprises (SOEs) have become increasingly important agents for fostering economic growth and developing functional capital markets in their countries of origin.

Chapter 3 focuses on the comparative return performance of the largest SOEs in Brazil, Russia, India and China (commonly known as the BRIC countries) against industry competitors in the US. In this paper we assume that the returns of the SOEs and the comparable US industries can be explained by the Fama and French three-factor model plus momentum (Carhart, 1997; Eugene F. Fama and French, 1993, 1998).

However, the aim of Chapter 3 is not to assess the average performance of SOEs but to evaluate their performance during crisis periods. If SOEs outperform comparable non-state-owned firms during a financial crisis the government's stake in the firm can be seen to act as a cushion for the state-owned enterprise, and thus protects the value of the firm to some degree. The economic reason for this "cushion" effect is that the stake of the government is so large that the firm is protected from general market conditions. We test for this cushion effect by analysing the sensitivities of the four-factor model during different phases of the GFC using a quantile regression framework.

1.3. Testing for differences in sovereign spreads during the GFC using propensity-matching estimators

In Chapter 4 we test for and measure contagion in sovereign debt markets using an approach that is more robust to exogenous crisis dating than standard approaches. We use propensity matching combined with an average treatment effect on the treated (ATET) method to correct possible sample selection effects on contagion tests. Propensity matching borrows from the methods of randomised controlled trials: at the first stage, general factor models, including crisis dummies, are fitted to the whole sample; then a set of non-crisis observations most closely matching the factor values of the crisis sample observations are drawn, building an artificial but matching "control" sample; and finally, the crisis and artificial non-crisis samples are compared in formal tests of shifts in spreads. Therefore, by allowing our crisis observations to act as "treated" units, we can test whether the difference in spreads versus our "non-treated" benchmark is statistically significant. We apply this method to test for contagion in sovereign debt markets during the recent crises.

1.4. The effects of the GFC on Colombian local-currency bond prices: an event study

In Chapter 5 we use high-frequency data on the Colombian local-currency bond market to measure the effects of the Global Financial Crisis (GFC). We assume that the US market acted as a transmission mechanism for the crisis in a standard market model. We also control for confounding effects by taking into account the effect of global, regional, and local macroeconomic surprises in the period before, during, and after the GFC. In order to model the effect of the surprises on bond returns we apply the method suggested by Balduzzi et al. (2001). We hypothesise that if local Colombian currency bonds decoupled during the GFC, average abnormal returns for Colombian bonds should have been negative or at least lower than during other comparable periods.

In Chapter 5 we provide an interesting set of related results such as the exact timestamp of the major events during the GFC and the effect of global, regional, and local macroeconomic surprises on Colombian local-currency bonds at the exact time of release. However, the most interesting finding relates to the size and sign of abnormal returns to local-currency Colombian bonds during the crisis.

Chapter 2 Can financial autarchy prevent contagion? The case of Colombian pension funds during the subprime, financial, and sovereign debt crises[1]

2.1 Introduction

The benefits and costs of regulation for financial crisis prevention and management have been intensely debated. A key issue is the extent to which regulators should attempt to quarantine domestic assets from external shocks (Binici et al., 2010; Houston et al., 2012). The question is especially important for retirement savings, where pension fund members may be compelled to contribute and invest according to regulation rather than at their own discretion, and account balances are preserved until a prescribed age. Financial contagion may be especially damaging to retirement welfare if negative returns occur late in working life or early in retirement when accumulations are highest. On the other hand, portfolio restrictions can create costly inefficiency. All of these influences are more severe in emerging economies, where members of pension funds have scant personal resources to buffer against financial shocks or inefficiency.

Here we study financial contagion originating from US equity markets during the subprime crisis, and its aftermath in the European sovereign debt crisis, as it affected regulated assets in emerging-economy pension funds. We divide the GFC into three contiguous episodes: the subprime, credit crunch (CCC), and European sovereign debt (ESD) crises. Using these three episodes, we test for evidence of contagion from source markets to a restricted portfolio, or autarchic asset, that is, Colombian private pension funds. We define autarchy as occurring when local

[1] Another version of this chapter has been published in the article by Edgardo Cayon and Susan Thorp titled: "Financial Autarchy as Contagion Prevention: The Case of Colombian Pension Funds", Emerging Markets Finance and Trade, May-June 2014, Vol. 50, Supplement 3, pp. 127-145.

investors have a limited choice in portfolio selection and ownership of the autarchic asset is legally mandatory. We define contagion as a significant change in co-movements of returns across markets, conditional on a crisis occurring in one market or group of markets. Contagion implies the creation of a new transmission channel above tranquil period conditions.[2] The transmission mechanism here is an idiosyncratic shock from a source asset market (e.g., from US equity markets during the subprime crisis), which transmits to other financial markets through the liquidation of international assets by investors. Investors liquidate to cover their losses rather than because of changes in fundamental valuations in the receiving market (Boyer et al., 2006; Dungey and Martin, 2007). Consequently, these actions by investors, usually in the country of origin, affect local markets in the receiving country and consequently the local investors.

The fact that countries are generally vulnerable to systemic crisis is not contentious; there is ample historical evidence that the causes of financial crises are not unique to each event (Reinhart and Rogoff, 2008). However, the recent financial crises originated in mature markets and were transmitted to emerging economies, rather than the previously more common reverse case (Fry et al., 2011). And although crises are, and will be, recurrent phenomena, there is some evidence that their effects could be mitigated by regulation.

Our contribution to the current body of literature will centre on the isolation or integration of emerging-market financial institutions, specifically, pension funds. Prior to 2008, some studies argued that the ability of emerging economies to withstand crises had improved (Felices and Wieladek, 2012; Kılınç et al., 2012;

[2] See, among others: (Bekaert et al., 2005; Dungey et al., 2005; Forbes and Rigobon, 2002; Pericoli and Sbracia, 2003). Some authors further refine this definition of contagion as correlation over and above economic fundamentals which are linked usually to macroeconomic indicators such as GDP growth, Balance of Payments or level of reserves (Bekaert et al., 2005; Boyson et al., 2010)

Köksal and Orhan, 2013; Powell and Martinez S., 2008). But despite reforms to financial institutions, stronger reserve positions, and restrictions on foreign exchange exposures, any decoupling of emerging economies from external economic and financial shocks appears to have been short-lived, especially in the face of the Great Recession (Dooley and Hutchison, 2009; Felices and Wieladek, 2012; Won et al., 2013). Latin America and Asia have been very exposed to macroeconomic trade factor shocks during the recent crisis (Bagliano and Morana, 2012).

Although there is general agreement that there has been no permanent decoupling of emerging markets from global shocks, there is evidence for temporary changes that are worth further exploration. Several studies find evidence of changing integration between emerging economies and developed economies. For example, Dufrénot et al. (2011) show that stock market volatility of Latin American economies with strong financial links to the US (such as Mexico and Chile) was highly sensitive to bad news from US banking and credit markets using data from January 2004 until April 2009. However, stock market volatility for Brazil, Colombia, and Peru, where access to US capital markets was partly limited by regulation, seemed more affected by changes in regional stock market volatility than offshore interest rates and credit spreads. And while multi-country studies find a generally higher degree of integration at the height of the crisis, correlation in the credit default swap markets increased more for developed than for emerging markets (Ping and Moore, 2008).

By analysing the extent of contagion to restricted pension assets in an emerging economy, we shed light on the effect of regulation on an important kind of inaccessible asset. We ask whether quarantining domestic savings can protect pension fund members from pervasive global financial shocks and under what circumstances this protection is effective.

We begin by describing some key features of private pension funds in Colombia, establishing that the funds are a financial autarchic asset. After describing our data, crisis episodes, and other descriptive statistics for the three crisis episodes, we set out contagion models and present a new dynamic measure of systematic risk. We then present estimation results and conclusions.

2.2 Background: pension funds

Latin American private pension funds have grown rapidly in the past few decades. By 2010, the approximate amount of assets under management in Latin American funds was estimated at around US$445 billion[3] and their annual rate of growth of assets under management has been as high as 25.4 per cent in recent years.

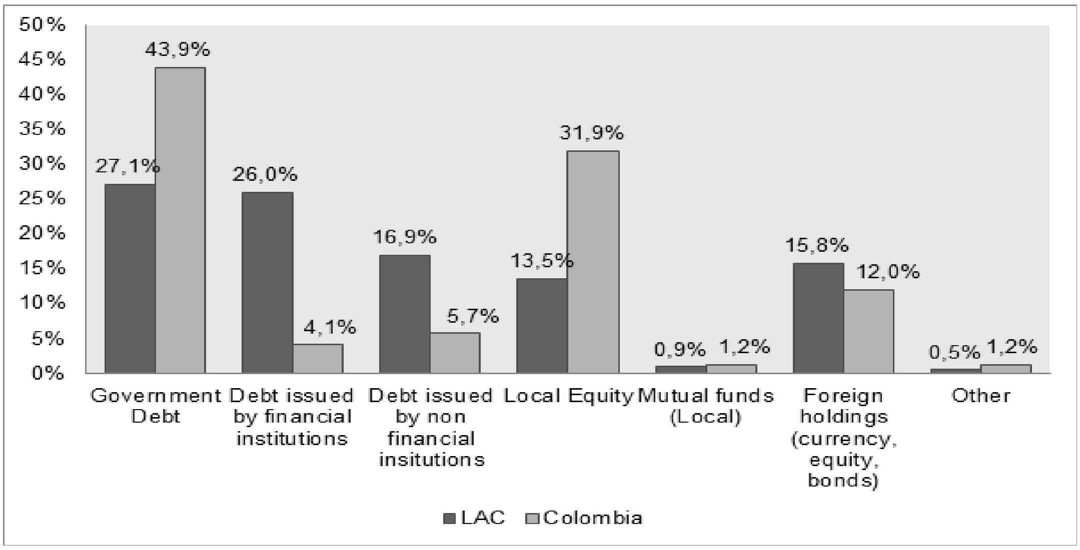
In the prevalent defined contribution (DC) pension plans, the employer and/or employee pay a fixed contribution to a savings fund, and the amount of pension eventually drawn by the employee depends on his/her level of savings and investment returns, net of administrative costs and taxes, at retirement. All the risk is borne by the employee with no further legal responsibilities for either the government or the employer (Impavido and Tower, 2009). Compared with defined benefit plans, DC plan providers may have an incentive to engage in riskier investments in order to compete with other participants in the market. Since the consequences of an unfavourable outcome can be disastrous to the members, and given the potential for agency problems in the investment process, governments tend

[3] Data retrieved from the Asociacion Internacional de Organismos de Supervison de Fondos de Pensiones (AIOS, 2011) http://www.aiosfp.org/estudios_publicaciones/estudios_pub_boletin_estadistico.shtml.

to set regulatory constraints on the types of assets that private funds can invest in (Arrau and Schmidt-Hebbel, 1995).

Figure 1 graphs the average investment limits by asset class in 2010 in the Latin American region, and for Colombia. One interesting feature of the asset allocations across Latin America is a preference for local-currency assets, on average 84.9 per cent of total investments as at 2010. Figure 2 graphs the local-currency assets in the Latin American region and shows that Colombian pension funds' local-currency holdings are above average for the region. This may ensure that local-currency pension liabilities (retirement incomes) are matched with assets while offering a cushion against volatile international capital that has triggered many past crises in the region.[4] Further, limiting holdings of offshore assets by local investors may help to increase the breadth and depth of local stock markets (AIOS, 2011).

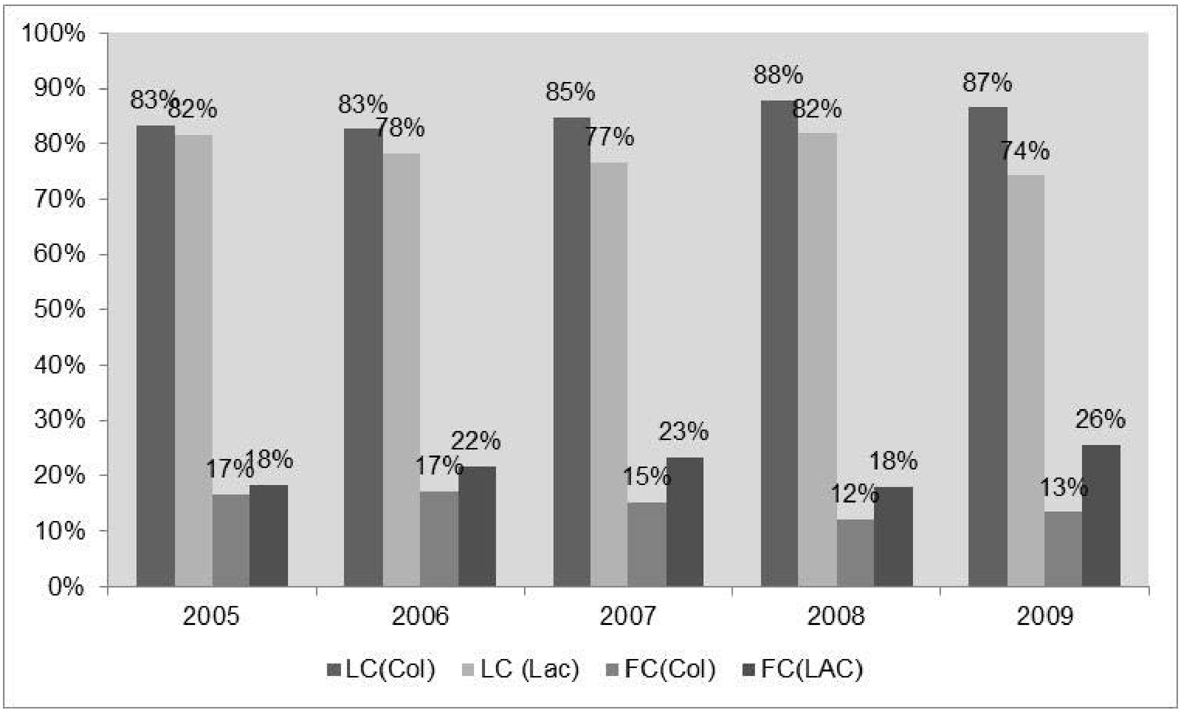
Figure 1: Average investment limits by asset class in private pension funds in Latin America countries (LAC) and Colombia, 2010



Source: (AIOS, 2011)

[4] The effects of capital flows as triggers for financial crisis in Latin America and its policy effects have been documented in the literature: Kaminsky et al. (2003); Kaminsky and Reinhart (1999); Calvo and Reinhart (1999); Edwards (1998) and more recently Dufrénot et al. (2011).

Figure 2: Evolution of foreign- and local-currency holdings of pension funds in Latin America countries (LAC) and Colombia (Col), 2005–2009.



Source: (AIOS, 2011)

From Table 1 we observe that Colombian pension portfolios have historically been concentrated in government and public entity debt at levels well above the region's average. Although regulation allows for a maximum of 31.9 per cent in local equity and 12 per cent in foreign holdings, this is often not reached: Colombian pension funds are on average 80 per cent in fixed-income instruments and 20 per cent in all other asset classes. As at August 2011, around 90 per cent of the investments were concentrated in domestic-currency fixed income and equities, and just 10 per cent of the investments were in foreign holdings.[5] The majority of domestic currency assets were sovereign and government entity debt (over 75 per

[5] In the short-term deposits in foreign currency there is also a negligible portion of hedging exchange derivatives, which amount to less than 0.005 per cent of the total investment composition (Superfinanciera, 2011). Colombia has a floating exchange regime and its currency is allowed to float freely against the US dollar, which is the reserve currency of choice in the Latin American region.

cent of total assets) with 10 per cent allocated to local equities and mutual funds (Superfinanciera, 2011).

Table 2.1: Evolution of investments by asset class in Colombian private pension funds (CPF) and for pension funds in the Latin American region (LAC), 2005–2009.

	2005	2006	2007	2008	2009
Government debt (Col)	47.31	47.19	44.12	48.35	42.00
Government debt (LAC)	43.73	38.96	34.07	41.14	34.81
Debt issued by financial institutions (Col)	9.19	8.05	7.71	9.74	5.12
Debt issued by financial institutions (LAC)	16.78	15.70	17.70	16.50	11.27
Debt issued by non-financial institutions (Col)	14.35	12.14	10.10	8.24	6.35
Debt issued by non-financial institutions (LAC)	9.40	9.25	9.11	10.68	11.27
Local equity (Col)	11.24	14.53	22.33	20.02	31.66
Local equity (LAC)	10.20	12.65	13.85	11.98	15.48
Mutual funds (Col)	1.18	0.77	0.40	1.52	1.41
Mutual funds (LAC)	1.47	1.77	1.87	1.55	1.47
Foreign holdings (currency, equity, bonds) (Col)	12.39	14.09	11.93	9.38	11.63
Foreign holdings (currency, equity, bonds) (LAC)	16.31	19.74	21.62	16.41	24.02
Other (Col)	4.34	3.23	3.41	2.75	1.83
Other (LAC)	2.11	1.94	1.78	1.74	1.70

Source: Data retrieved from the Asociacion Internacional de Organismos de Supervison de Fondos de Pensiones (AIOS, 2011) . (Col) stands for Colombia and (LAC) for Latin American countries.

The effect of these regulatory constraints has become more evident during the subprime crisis. For example, Pino and Yermo (2010) observed that the real average annual rate of return for private pension funds in OECD countries in the year 2008 was -24.1 per cent, with this large loss blamed on exposures to equity during the early stages of the crisis. By contrast, some funds from non-OECD countries (including Colombia) did not suffer losses at all during 2008 due to their high exposure to local government debt.

High exposure to local currency assets and regulatory restrictions make Colombian pension funds an interesting case study in the effects of regulation in time of crisis. First, Colombia is a conservative benchmark for other pension systems in the Latin American region. Secondly, daily price data for the Colombian funds are publicly available, which allows us to observe the transmission of an idiosyncratic shock from the US and its effect on the Colombian pension fund returns on a high-frequency basis.

Investments in Colombian private pension funds can be labelled as autarchic financial assets where autarchy is obtained through a restrictive regulatory framework that favours low-risk investments in local currency. If regulation can act as a cushion for cash flow volatility in times of financial crisis, then it may be possible to reduce contagion effects. However, one important caveat is that by imposing the quantitative restrictions, pension fund managers are limited to a constrained portfolio that may not grasp the full benefits of diversification (Davis, 2001), possibly generating a suboptimal risk and return over the long run. Here we restrict ourselves to assessing the extent to which these restrictions can reduce contagion during crises, as a necessary, but not sufficient, condition for justifying regulation.

2.3 Data and summary statistics

The pension fund data for the present study have been retrieved from ASOFONDOS,[6] the private pension funds association in Colombia. Our sample

[6] The data was retrieved from *El Centro de Informacion Consolidada Asofondos* website (Asofondos, 2011). In Colombia, the regulation concerning the private pension system and its investment regime is contained in the *ley 100 de 1993*, as drafted and approved by the Colombian Congress. All subsequent

contains the daily net asset value (NAV) per unit for each provider (as required by Colombian regulation) from February 1, 2005, to August 31, 2011. Before 2005, the Colombian private pension funds (CPF) were required by law to pay a guaranteed minimum monthly rate of return, which was usually a few basis points higher than the inflation rate reported by the government. For this reason we omit the period before February 2005. The minimum return requirement was later modified by the *Decreto 1592 de 2004*, which created a benchmark index that consisted of a weighted average of the reported annual returns of all Colombian private pension funds, the Colombian equity market index, and an international benchmark equity index (the US S&P 500), scaled down by 70 per cent, and adjusted for the amount of local equity or international investments held in each PPFs portfolio.[7] This regulation allowed the Colombian pension funds to allocate their investments across the financial instruments sanctioned by the *Superintendencia Financiera de Colombia* (SFC) investment regime.[8] (The operational regulatory framework which includes the mandatory investment regime for CPF operators as well as the guidelines for calculating the NAV for each fund is contained in a series of documents issued by the SFC.)[9]

For the period under study, we measure composite fund performance by adding each of the funds' reported NAV per unit for a given day on an equally weighted basis. Summary statistics for each pension fund and the composite

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regulations regarding the day-to-day operations of the private fund industry in Colombia are the responsibility of the *Superintendencia Financiera de Colombia (SFC),* which is the agency responsible for overseeing and regulating the activities of the financial sector in Colombia and a part of the *Ministerio de Hacienda* (Ministry of Economics and Finance).

[7] The formula for the minimum rate of return is available in Appendix A.

[8] A comprehensive list of the financial instruments can be found in the relevant articles of the *Decreto 2555 de 2011*, which collects all previous regulations on the matter available at http://www.asofondos.org.co/VBeContent/newsdetail.asp?id=19&idcompany=3

[9] These documents are: Circular Externa 007 de 1996, Circular Externa 036 de 2003, and the Decreto 1592 de 2004, Decreto 2175 de 2007, Decreto 4935 de 2009, and Decreto 2555 de 2011, which can be found at the SFC website: http://www.superfinanciera.gov.co; the NAV calculation methodology can be found in Appendix B.

performance measure are presented in Table 2.2. Pension fund returns are negatively skewed with a high kurtosis, a fact guiding our choice of volatility model for contagion measurement.

Table 2.2: Descriptive statistics: annualised daily returns to Colombian private pension funds (CPF), February 2, 2005, to August 31, 2011.

	CitiColfondos	Horizonte	Porvenir	Protección	ING	CPF
Mean	8.29	9.38	7.97	9.87	8.44	8.80
Median	7.75	9.84	8.75	8.78	8.19	8.61
Maximum	22.48	21.98	18.9	25.04	21.07	21.35
Minimum	-3.10	-1.88	-3.57	-2.83	-4.10	-3.02
Std. dev.	5.88	5.82	5.79	6.20	6.03	5.83
Skewness	-0.75	-0.62	-0.39	-0.71	-0.67	-0.71
Kurtosis	8.60	8.83	8.64	8.98	9.69	8.73
Jarque-Bera	2402	2535	2315	2696	3322	2490
Probability	0.00	0.00	0.00	0.00	0.00	0.00
Observations	1714	1714	1714	1714	1714	1714
Members* ('000)	1,533	1,701	3,096	2,050	1,161	9,541
Assets USD mil*	7,045	8,077	15,074	12,563	6,508	49,266

*As reported by Superfinanciera in January 2012

Table shows summary statistics for daily percentage changes (annualised) in net asset values per unit of Colombian pension Funds, February 2, 2005, to August 31, 2011. CPF is calculated by adding all the funds' reported net asset value per unit each day on an equally weighted basis.

We use daily returns to US stocks (the S&P 500) as the common factor and source of financial shocks over the same sample. For robustness, we also repeat the same tests using the J.P. Morgan Emerging Market Bond Index (EMBI)[10] and the Eurozone government debt index (EFFA)[11] as sources of shocks. A regional equity portfolio is proxied by the MCSI Emerging Market Latin American (*LAC*)[12] index, and the Colombian Peso-USD exchange rate (*COPFX*) is used to measure the contribution of the floating exchange regime to the CPF funds' performance. These series were collected from Bloomberg.

We divide the sample into a pre-crisis phase followed by three contiguous crisis phases. The first phase subprime crisis begins July 26, 2007, which was the day the Dow Jones recorded a significant large loss in response to bad news from mortgage lender Countrywide Financial. At this point, the market processed news of "difficult conditions" in the subprime market following Countrywide Financial Corporation's SEC filing on July 24. The beginning of the credit crunch crisis (CCC) is generally dated from the time Lehman Brothers filed for bankruptcy on September 15, 2008. The European sovereign debt crisis (ESD) we date from October 22, 2009, when Fitch first downgraded and reported a negative outlook for Greek sovereign debt. The turbulence in European debt markets was continuing at the end of our

[10] The Emerging Market Bond Index Global is a market-weighted capitalisation index of USD-denominated bonds and Eurobonds from sovereign and quasi-sovereign entities and is a worldwide recognized benchmark for emerging-markets debt.
[11] The EFFA/Bloomberg index, which includes all of the Eurozone government debt with maturities of more than one year, is the most comprehensive Eurobond index, including more than 364 issues from all members. The EFFA is a market-weighted capitalisation index.
[12] The MSCI EM (Emerging Markets) Latin America Index is a free float-adjusted market capitalisation weighted index that is designed to measure the equity-market performance of emerging markets in Latin America. The MSCI EM Latin America Index consists of the following five emerging-market country indices: Brazil, Chile, Colombia, Mexico, and recently Peru, as of May 30, 2011.

sample, August 31, 2011,[13] which is the latest period for which we have Colombian pension fund data.

Since contagion is here defined as a significant change in co-movements of returns across markets, conditional on a crisis occurring in one market or group of markets, we begin by considering the dynamics of returns and co-movement with the S&P 500 as the common factor across the phases. Table 2.3 reports summary statistics for returns to Colombian pension funds (*CPF*), the regional equities index (*LAC*) and the Colombian peso/USD exchange rate (*COPFX*) along with co-movement with the S&P 500 for the pre-crisis, crises and total sample periods.

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[13] The key dates for the subprime and CCC crisis were taken from the financial turmoil timeline chart from the Federal Reserve Bank of St. Louis (http://timeline.stlouisfed.org/pdf/CrisisTimeline.pdf), and for the European sovereign debt crisis from the credit rating function in Bloomberg. There are other studies that use similar dates for the CCC and place the subprime around the same period, including Frank and Hesse (2009), Dooley and Hutchison (2009), and Felices and Wieladek. (2012).

Table 2.3: Descriptive statistics: sub-sample

	Colombian Pension Fund (CPF)	***S&P 500***	***LAC***	***COPFX***
		Pre-crisis		
Mean (% p.a.)	7.13	9.38	38.85	-8.24
St. Dev. (% p.a.)	3.40	10.34	23.12	8.44
Sharpe ratio	0.57	-0.29	-0.26	0.01
Beta (S&P)	0.08***	-	1.44***	-0.19***
		Subprime		
Mean (% p.a.)	5.75	-16.37	-8.35	4.87
St. dev. (% p.a.)	2.81	20.75	33.75	15.08
Sharpe ratio	0.35	-0.68	-0.26	0.01
Beta (S&P)	0.03***	-	0.94***	-0.10***
Mean ratio	0.81	-1.75	-0.21	-0.59
SD ratio	0.82	2.01	1.46	1.79
		CCC crisis		
Mean (% p.a.)	9.96	-14.63	1.14	-0.46
St. dev. (% p.a.)	3.39	33.17	47.61	18.10
Sharpe ratio	1.62	-0.58	-0.09	-0.27
Beta (S&P)	0.31***	-	1.03***	-0.12***
Mean ratio	1.40	-1.56	0.03	0.06
SD ratio	0.99	3.21	2.06	2.15
		ESD crisis		
Mean (% p.a.)	8.89	-5.17	0.89	-2.04
St. dev. (% p.a.)	4.28	27.59	38.78	14.87
Sharpe ratio	1.06	-0.42	-0.16	-0.43
Beta (S&P)	0.05***	-	1.01***	-0.13***
Mean ratio	1.25	-0.55	0.02	0.25
SD ratio	1.26	2.67	1.68	1.76

Table reports descriptive statistics for daily returns in local currency during pre-crisis period and crisis phases. An appreciation of the Colombian peso produces a negative return to COPFX. The Sharpe ratio is defined as $SR = \frac{(\overline{R_j} - \overline{R_f})}{\sigma_j}$ where $\overline{R_j}$= the mean return of the index over the sample, $\overline{R_f}$= the mean return of the risk-free rate, the US 10-Year Treasury index adjusted by the Colombian country premium, and σ_j= the standard deviation of returns to the index. The SD ratios and mean ratios are calculated using the sample statistics during the crisis period in the numerator and the pre-crisis period statistics in the denominator. A ratio greater than one signals an increase in volatility/return during the crisis period over the pre-crisis period. Significant coefficients at the *90%, **95%, ***99% confidence level.

Simple betas with S&P 500 returns for each index are reported in Table 2.3. For the pension funds (CPF) the beta declines from the pre-crisis levels during the subprime crisis, but is markedly higher during the ESD. Regional equity index (LAC) returns follow a similar pattern with a sharper increase in the post-Lehman (CCC) phase. The beta coefficient between S&P 500 returns and the exchange rate (COPFX) declines from the pre-crisis level during the subprime crisis but steadily increases in the next phases. The relatively low responsiveness of CPF returns to the first two crisis phases is corroborated by the standard deviation (SD) ratios, which are below one until the ESD crisis. The SD ratios and mean ratios are calculated using the sample statistic during the crisis period in the numerator and the pre-crisis period statistic in the denominator. A ratio greater than one signals an increase in volatility of returns during the crisis period over the pre-crisis period. For other returns series (S&P 500, *LAC* and *COPFX*) the SD ratio is greater than one for all crisis phases. Also, the Sharpe Ratio[14] for CPF tends to outperform all other investments during all the crisis periods under observation. Sharpe ratios for all indices are calculated from the perspective of the Colombian investor so that the returns are adjusted for the exchange rate, resulting in negative values in several cases. Other statistics are calculated in local currency.

Even though quantitative restrictions such as asset allocation caps and limited foreign currency exposure can limit the upward risk/return potential of CPF returns by constraining investment portfolios, these same measures also can limit the

[14] The Sharpe ratio is the most common performance investment measure and is defined as $SR = \frac{(\overline{R_j}-\overline{R_f})}{\sigma_j}$ where $\overline{R_j}$= the average return of the security, $\overline{R_f}$= the average return of the proxy of the risk-free rate, which in our case is the US 10-Year Treasury index adjusted by the Colombian country premium and σ_j= the standard deviation of the security.

potential for losses as observed by higher Sharpe ratios during the different episodes of the crisis. Simple sample statistics in Table 2.2 show up this trade-off.

2.4 Contagion model

We evaluate evidence for contagion via changes in the time-varying volatility of the pension fund returns across different crisis phases. Specifically, we search for significant changes in the proportion of the filtered returns volatility that can be attributed to US stock market shocks. For comparison, we conduct the same tests for the regional stock index and the exchange rate and then check for robustness using fixed income rather than the equity index as a source of shocks. By implementing and estimating a conditionally heteroskedastic model of daily returns volatility via a multivariate GJR-GARCH model (Glosten et al., 1993), we avoid the criticism of unconditional correlation comparisons noted by Forbes and Rigobon (2002).[15]

We begin by allowing the systematic shock to follow a process in the following form:

$$\begin{aligned} r_{\mathrm{US},t} &= \theta_{0,US} + \theta_{1,US} r_{\mathrm{US},t-1} + \varepsilon_{\mathrm{US},t} \\ r_{\mathrm{i},t} &= \theta_{0,i} + \theta_{1,i} r_{\mathrm{i},t-1} + \phi_1 r_{US,t} + \phi_2 r_{\mathrm{US},t} d_{1,t} + \phi_3 r_{\mathrm{US},t} d_{2,t} + \phi_4 r_{\mathrm{US},t} d_{3,t} + \varepsilon_{\mathrm{i},t} \end{aligned} \quad (2.1)$$

Where $r_{US,t}$ are daily returns to the S&P 500 and $r_{i,t}$ $i = 1, \ldots, 3$ are daily returns to *CPF, LAC, or COPFX* and $d_{j,t}$ $, j = 1, \ldots, 3$ are indicator variables taking the value 1 during the relevant crisis period and 0 otherwise. This model of the mean ensures that the residuals for the DCC-GJR GARCH model represent idiosyncratic risk from each market and have a zero mean and serially uncorrelated. The shocks are denoted as $\varepsilon_{us,t}$ and $\varepsilon_{i,t}$. Further, we treat the S&P 500 as the originating source of potential contagion but also allow the intensity of the impact of the common factor to shift

[15] Similar structures are estimated by Fujii (2005), Chiang et al. (2007), and Ping and Moore (2008).

dynamically. We do not include a separate time-varying regional factor as do Bekaert et al. (2005), since the Colombian pension funds are heavily biased towards domestic assets with little regional exposure, and we concentrate on testing for contagion sourced in the US markets directly into the domestic asset markets. However, we do check for misspecification and the influence of regional and global debt shocks by re-testing the model, treating shocks from those markets as possible sources.

We are primarily interested in changes in the conditional correlation between shocks from equation (2.1), once dynamic risk loadings are accounted for. The covariance of the risk of the S&P 500 with *CPF, LAC,* and *COPFX* respectively is:

$$h_{US,t} = \alpha_{0,US} + \alpha_{1,US}\varepsilon_{US,t-1}^2 + \alpha_{2,US}I_{US,t-1}\varepsilon_{US,t-1}^2 + \alpha_{3,US}h_{US,t-1}$$

$$h_{i,t} = \alpha_{0,i} + \alpha_{1,i}\varepsilon_{i,t-1}^2 + \alpha_{2,i}I_{i,t-1}\varepsilon_{i,t-1}^2 + \alpha_{3,i}h_{i,t-1}$$

$$h_{USi,t} = \alpha_{0,USi} + \alpha_{1,USi}\varepsilon_{US,t-1}\,\varepsilon_{i,t-1} + \alpha_{2,USi}I_{US,t-1}\varepsilon_{US,t-1}\,I_{i,t-1}\varepsilon_{i,t-1} + \alpha_{3,USi}h_{USi,t-1}$$

$$\begin{bmatrix}\varepsilon_{US,t}\\ \varepsilon_{i,t}\end{bmatrix}|\Omega_{t-1}\sim N\left(\begin{bmatrix}0\\0\end{bmatrix},\begin{bmatrix}h_{US,t} & h_{USi,t}\\ h_{USi,t} & h_{i,t}\end{bmatrix}\right), \tag{2.2}$$

where $h_{US,t}$ is the conditional variance of filtered returns from the US (origination country), $h_{i,t}$ is the conditional variance of the idiosyncratic risk of the local or regional market index under scrutiny, $h_{USi,t}$ is the covariance between the US market and the local or regional market index and $I_{j,t-1}$ is an indicator equal to one when $\varepsilon_{j,t-1}$ is negative and zero otherwise. In order to ensure that the relevant parameters are positive and avoid negative volatilities we used a diagonal BEKK specification.

Using fitted values from the model, we can compute a single index factor as:

$$\beta_{i,t} = \frac{h_{USi,t}}{h_{US,t}} \tag{2.3}$$

The advantage of this specification is that it further allows us to decompose the variance of the local index returns into systematic and idiosyncratic components:

$$h_{i,t} = {\beta_{i,t}}^2 h_{US,t} + h_{ei,t}, \qquad (2.4)$$

where $h_{i,t}$ is the variance of the index under scrutiny, ${\beta_{i,t}}^2 h_{US,t}$ is the part of the variance attributed to a common systematic transmission mechanism, and $h_{ei,t}$ is the part of the variance attributed to idiosyncratic factors. Therefore, expressing the variance decomposition as a proportion of systematic and idiosyncratic risk is straightforward:

$$u_{sys,i,t} = \frac{{\beta_{i,t}}^2 h_{US,t}}{h_{i,t}}$$

$$u_{idio,i,t} = \frac{h_{ei,t}}{h_{i,t}}$$

(2.5)

Evaluation of contagion from US stock market shocks proceeds along similar lines to Bekaert et al. (2005), by testing for breaks in the proportion of the residual variance due to systematic factors during each crisis phase. We regress the extracted systematic factors on a constant, a lagged value to control for serial correlation and indicator variables for the phases of the crisis:

$$u_{sys,i,t} = \varphi_0 + \varphi_1 u_{sys,t-1} + \varphi_2 I_{subprime,t} + \varphi_3 I_{CCC,t} + \varphi_4 I_{ESD,t} + e_{i,t}\text{ , or}$$

$$\beta_{i,t} = \gamma_0 + \gamma_1 \beta_{i,t-1} + \gamma_2 I_{subprime,t} + \gamma_3 I_{CCC,t} + \gamma_4 I_{ESD,t} + \epsilon_{i,t}$$

(2.6)

where $I_{j,t}$ is an indicator variable that take the value of one for the respective crisis dates (subprime, CCC, and ESD) and is zero otherwise. Contagion, in the form of changes to correlation in volatility transmissions from the systematic factor, is detected when coefficients on the crisis phase indicators are significantly different *from zero.*

2.5 Results

First, we find that in the case of pension fund returns, all coefficients on the crisis-period indicator variables (ϕ_i) are significantly greater than zero, denoting dynamic changes in the risk loadings between the S&P 500 and CPF. Estimates of the GJR-GARCH show that the coefficient on $I_{j,t-1}$ is significant and positive, so that negative shocks have a greater effect than positive shocks on volatility.[16] In other words, the correlation between the US equity returns and CPF returns intensified during the crisis periods. However, we are interested in testing further for contagion from the shocks from this model, after accounting for volatility dynamics. The graphs of the dynamic path of $u_{sys,i,t}$ and $u_{idio,i,t}$ highlight the changing impact of US stock market shocks on the inaccessible Colombian pension funds returns (*CPF*), the more tradable regional stock index (*LAC*), and the exchange rate (*COPFX*). Figure 3 graphs the decomposition of the conditional volatility of Colombian pension fund returns into systematic (US-sourced) and idiosyncratic components.

[16] Additionally, we allowed for the interaction of dummies representing positive and negative returns with our estimate of systematic risk (equation 2.6). We found that negative shocks have a slightly greater effect in systematic risk than positive ones in the case of LAC and COPFX and the opposite holds in the case of CPF except when EMBI is chosen as the factor, although the differences in size are not large. We also allowed for this interaction to be crisis specific and obtained similar results.

Figure 3: Conditional variance decomposition, Colombian pension fund (CPF) returns

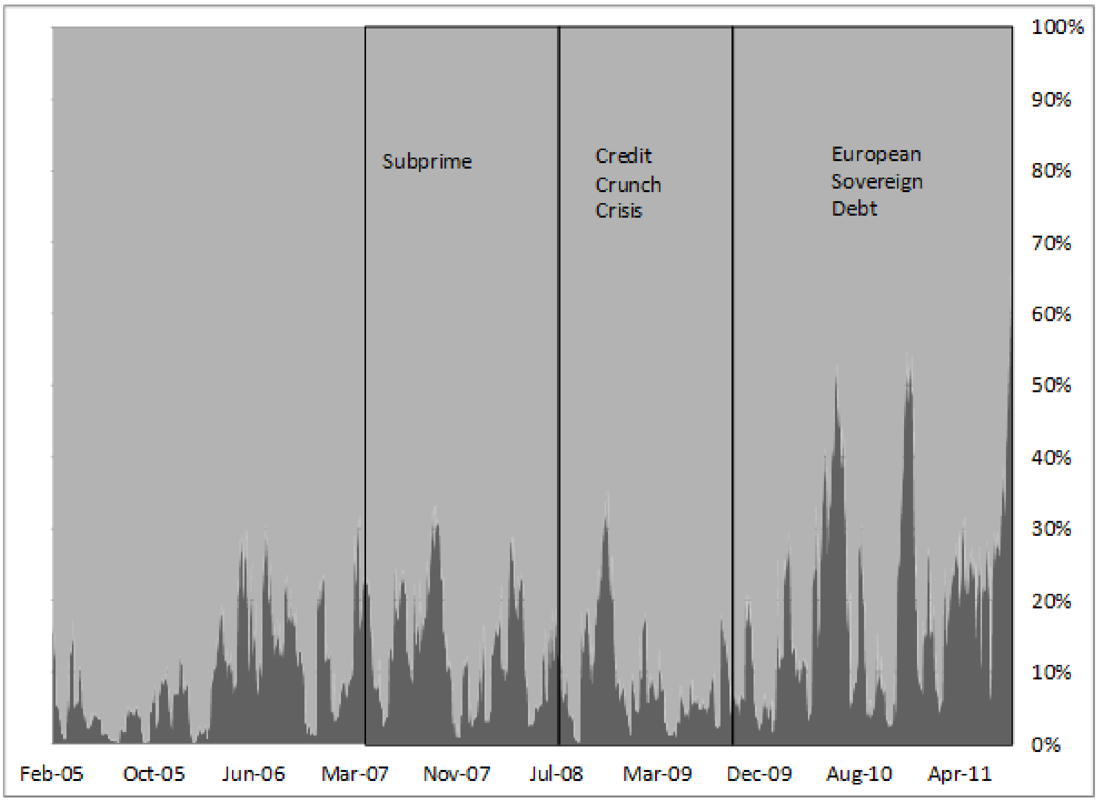

Figure 3 graphs proportion of conditional volatility of returns to Colombian pension funds due to systematic volatility shocks from US S&P 500 (dark grey).

Systematic risk increases at the beginning of the subprime crisis, decreases during the CCC and then increases dramatically during the ESD crisis. It appears that initial responses to the subprime and CCC crises dissipate as markets focus more on the fundamentals affecting pension fund performance, as opposed to the ESD, in which the effects of bad news in fixed-interest markets show up as persistently higher systematic variance.

Figure 4, which decomposes the conditional variance of *LAC*, shows that the total average systematic risk observed for the whole sample is higher than that observed for CPF. Consistent with some evidence for decoupling/recoupling, there is an increase in systematic risk before the subprime crisis that dampens during the crisis then escalates during the CCC and ESD periods.

Figure 4: Conditional variance decomposition, Colombian pension fund (CPF) returns

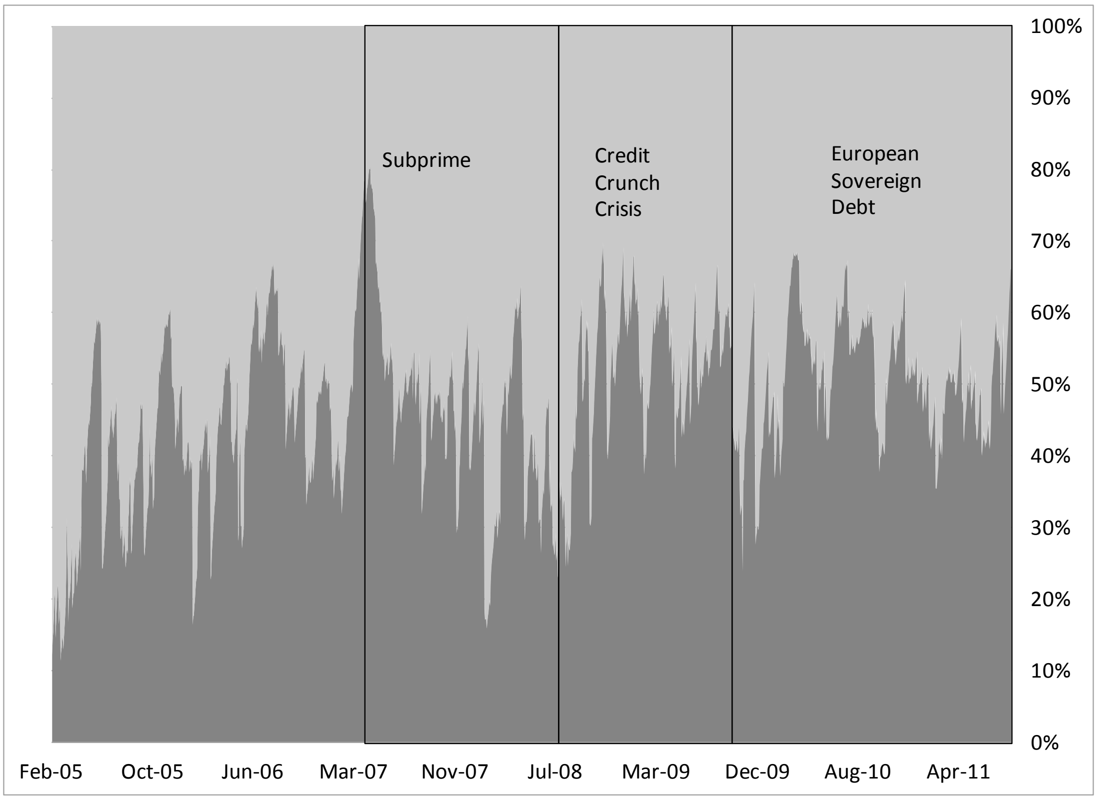

Figure 4 graphs proportion of conditional volatility of returns to the Latin American regional stock index due to systematic volatility shocks from US S&P 500 (dark grey).

On the other hand, *COPFX* systematic risk in Figure 5 shows a dampening effect during the subprime crisis and a sudden increase towards the end of the CCC, to levels which are maintained throughout the ESD crisis. Finally, the increase in systematic risk during mid-2006 for both CPF and *COPFX* is explained by three consecutive interest hikes by the Federal Reserve between March and June of 2006 that had strong effects in Latin American markets (Ocampo, 2009). Three rapid hikes caused turbulence, as the cost of financing adjusted to new levels, and capital flows to the region diminished, while foreign borrowers reassessed their investments.

Figure 5: Conditional variance decomposition, Colombian pension fund (CPF) returns

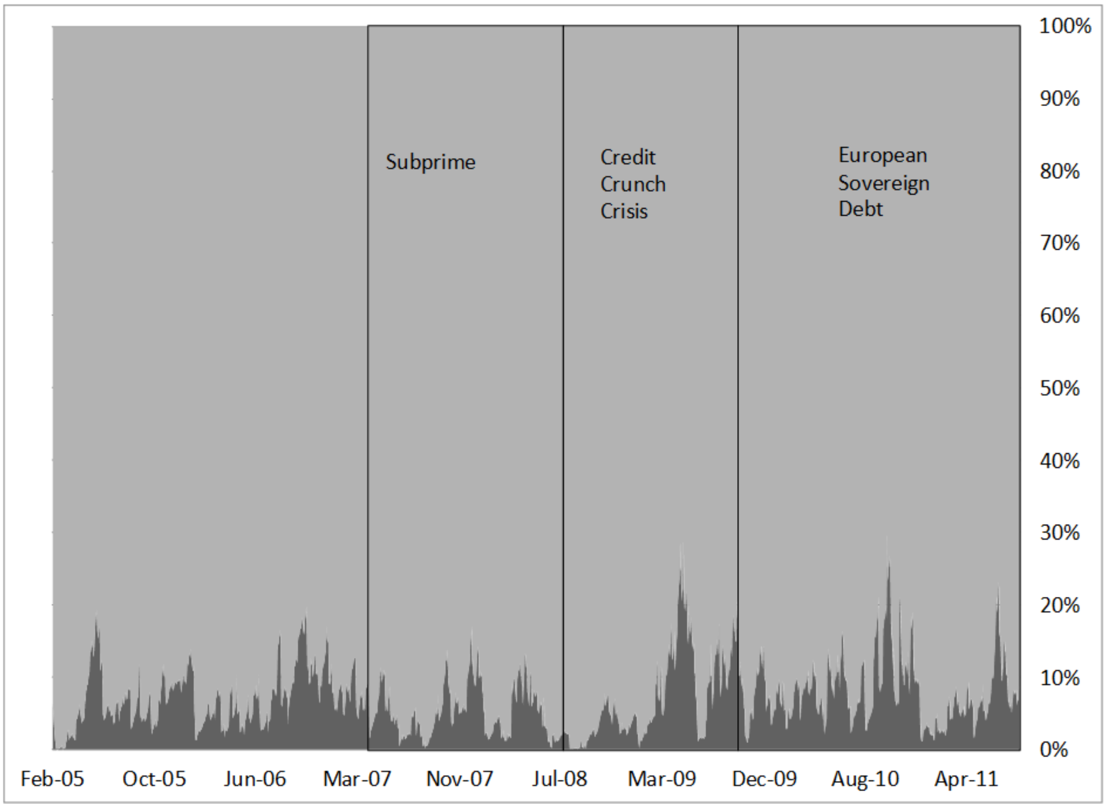

Figure 5 graphs proportion of conditional volatility of returns to the Colombian peso/USD exchange rate due to systematic volatility shocks from US S&P 500 (dark grey).

2.6 Contagion from US stock market factor

In order to investigate the presence of contagion, we conducted the test described in equation (2.6) for each of the systematic variance components $u_{sys,i,t}$ and time-varying factor $\beta_{i,t}$ respectively. The results are summarised in Table 2.4.

Table 2.4: Tests for contagion: US stock market to Colombian pension fund (CPF), LAC stocks and USD/Colombian peso exchange rate

Dependent Variable							
	$u_{sys,i,t}$				$\beta_{i,t}$		
Coefficient	*CPF*	*LAC*	*COPFX*	*Coefficient*	*CPF*	*LAC*	*COPFX*
$\varphi_{2Subprime}$	0.001	-0.001	0.000	$\gamma_{2Subprime}$	-0.001	-0.016***	0.002
	[0.002]	[0.002]	[0.001]		[0.001]	[0.006]	[0.002]
φ_{3CCC}	0.001	0.003	0.002	γ_{3CCC}	-0.001	-0.020***	-0.001
	[0.002]	[0.002]	[0.001]		[0.001]	[0.006]	[0.002]
φ_{4ESD}	0.005***	0.003	0.002	γ_{4ESD}	0.002*	-0.013***	0.000
	[0.002]	[0.007]	[0.001]		[0.001]	[0.007]	[0.002]
R-squared	0.922	0.925	0.859	R-squared	0.960	0.930	0.896
S.E. of regression	0.032	0.032	0.019	S.E. of regression	0.014	0.078	0.032

Table 2.4 reports tests for contagion from US stock market returns to Colombian pension fund returns (*CPF*), Latin American stock index returns (*LAC*) and the USD/Colombian peso exchange rate (*COPFX*) over three crisis phases. The LHS of the table reports coefficient estimates obtained from regressing the proportion of conditional variance of returns to each local index due to US stock market shocks on a constant and indicators for crisis phases, $u_{sys,i,t} = \varphi_0 + \varphi_1 u_{sys,t-1} + \varphi_2 I_{subprime,t} + \varphi_3 I_{CCC,t} + \varphi_4 I_{ESD,t} + e_{i,t}$. The RHS of the table reports coefficient estimates obtained from regressing the single index beta (US stock market shocks) on a constant and indicators for crisis phases, $\beta_{i,t} = \gamma_0 + \gamma_1 \beta_{i,t-1} + \gamma_2 I_{subprime,t} + \gamma_3 I_{CCC,t} + \gamma_4 I_{ESD,t} + \epsilon_{i,t}$. $\varphi_5 \text{POV} * (u_{sys} * I < 0)$ in which I=1 if return in CPF, LAC or COPFX is positive and 0 otherwise and . $\varphi_5 \text{NEG} * (u_{sys} * I < 0)$ in which I=1 if return in CPF, LAC, or COPFX is negative and 0 otherwise. Significant coefficients indicate contagion at the *90%, **95%, ***99% confidence level.

The only significant evidence of contagion from US stock volatility to the inaccessible asset, *CPF,* relates to the European sovereign debt crisis period. No significant volatility contagion is evident in the earlier periods despite the severity of the shocks. Like many other contagion studies (e.g., Dungey et al., 2010) we find some evidence of weakening links between US and regional LAC volatility during the crisis phases. Overall, graphical evidence and structural break tests confirm that despite intensified risk factors, the regulated assets were protected from stock market volatility contagion during the subprime crisis and its aftermath in the post-Lehman credit crunch, but that shocks from fixed interest markets may still be transmitted, despite limited international exposure.

2.7 Contagion from bond market factors

We drill down into the sensitivity of the autarchic assets to bond market shocks using fixed-income indices as the potential sources of contagion. There are large investments in Colombian local government debt in CPF, so bond market shocks may be more relevant. The results are summarised in Table 2.5.

Table 2.5: Tests for contagion: emerging-market bond index and European bond index to Colombian pension funds (CPF)

Dependent Variable					
	$u_{sys,CPF,t}$			$\beta_{CPF,t}$	
Coefficient	*EMBI*	*EFFA*	*Coefficient*	*EMBI*	*EFFA*
$\varphi_{2Subprime}$	-0.006	0.004***	$\gamma_{2Subprime}$	-0.017**	-0.01**
	[0.004]	[0.001]		[0.007]	[0.005]
$\varphi_{3\mathrm{CCC}}$	0.003	0.001	$\gamma_{3\mathrm{CCC}}$	-0.011	-0.006
	[0.004]	[0.001]		[0.007]	[0.005]
φ_{4ESD}	0.000	0.003***	γ_{4ESD}	0.016***	-0.013**
	[0.003]	[0.001]		[0.006]	[0.005]
R-squared	0.848	0.832	R-squared	0.847	0.852
S.E. of regression	0.054	0.018	S.E. of regression	0.099	0.072

Table 5 reports tests for contagion from the emerging-market bond index return (EMBI) and European bond index returns (EFFA) to Colombian pension fund returns (*CPF*) over three crisis phases. The LHS of the table reports coefficient estimates obtained from regressing the proportion of conditional variance of *CPF* returns due to each bond index shock on a constant and indicators for crisis phases, $u_{sys,i,t} = \varphi_0 + \varphi_1 u_{sys,t-1} + \varphi_2 I_{subprime,t} + \varphi_3 I_{CCC,t} + \varphi_4 I_{ESD,t} + e_{i,t}$. The RHS of the table reports coefficient estimates obtained from regressing the single index beta (index return shocks) on a constant and indicators for crisis phases, $\beta_{i,t} = \gamma_0 + \gamma_1\beta_{i,t-1} + \gamma_2 I_{subprime,t} + \gamma_3 I_{CCC,t} + \gamma_4 I_{ESD,t} + \epsilon_{i,t}$. φ_5POV $*$ $(u_{sys} * I < 0)$ in which *I*=1 if return in CPF is positive and 0 otherwise and . φ_5NEG $*$ $(u_{sys} * I < 0)$ in which *I*=1 if return in CPF is negative and 0 otherwise. Significant coefficients indicate contagion at the *90%, **95%, ***99% confidence level.

Some tests of $\beta_{CPF,t}$ support a finding of contagion from the EMBI during the subprime and ESD crises but not the CCC crisis. However, there is no evidence of contagion from the EMBI in any of the crises in regressions of $u_{sys,CPF,t}$. For the European index, EFFA, we find that there is evidence of contagion in the subprime and ESD crisis. This is an intriguing result since there are two reasons why one would expect the EMBI to be more significant than the EFFA as a transmission factor for contagion in CPF. First, the EMBI is computed from emerging-market issuers (Greece included) with similar weights, and Colombia is also a small component of the index. Secondly, the EFFA is a market-weighted index for the whole Eurozone, so the effect of volatility of the countries in crisis should be mitigated by the larger and more stable countries in the index. However, since the majority of significant contagion is related to EFFA volatility, we conclude that it is a better proxy for the systematic factor than the S&P 500 in the case of CPF.

2.8 Quantile regression

One feature of the volatility models implemented here are asymmetries in the distributions of statistics of interest, in particular the conditional systematic variation proportion, $u_{sys,i,t}$. We implement quantile autoregression QAR(1) as proposed by Koenker and Xiao (2006) in order to check the robustness of contagion tests (see Appendix C). We generalise equation (2.6) and allow coefficients to be quantile dependent,

$$F^{-1}(\tau|u_{sys,i,t}) = \varphi_0(\tau) + \varphi_1(\tau)F^{-1}(\tau|u_{sys,i,t-1}) + \varphi_2(\tau)I_{subp,t} + \varphi_3(\tau)I_{CCC,t} + \varphi_4(\tau)I_{ESD,t} + e_{i,t} \quad (2.7)$$

where τ represents the 1%, 5%, 10%, 25%, 50%, 75%, 90%, 95%, and 99% quantiles respectively. By estimating over the extreme quantiles we can observe the structure of dependence while taking into account the asymmetric nature of our data (Baur, 2012). Specifically, we are interested in whether unusually large systematic volatility proportions are more common during crisis phases. Table 2.6 outlines the results. Consistent with earlier results and Figure 2, contagion from US stocks to pension funds is primarily confined to extreme volatility events during the subprime and sovereign debt crisis.

Table 2.6: Tests for contagion by quantile: US stock market to Colombian pension fund (CPF), LAC stocks and USD/Colombian peso exchange rate

$u_{sys,i,t}$				
	Quantile	***CPF***	***LAC***	***COPFX***
$\varphi_{2\,Subprime}$	1%	0.001	0.006	0.002
	2%	0.002	-0.010	0.000
	5%	-0.004	-0.015	-0.002
	10%	-0.002	-0.006	-0.002
	90%	0.005**	0.003	-0.001
	95%	0.019*	0.009	0.003
	98%	0.027**	0.001	0.006
	99%	0.018*	-0.007	0.028*
	Quantile	***CPF***	***LAC***	***COPFX***
$\varphi_{3\,\mathrm{CCC}}$	1%	0.001	0.067**	0.001
	2%	0.004	0.020	-0.001
	5%	-0.003	-0.021	-0.001
	10%	0.001	-0.002	0.000
	90%	0.006	0.003	0.007
	95%	0.006	0.000	0.020**
	98%	0.016	0.000	0.033***
	99%	0.024	-0.003	0.028**
	Quantile	***CPF***	***LAC***	***COPFX***
$\varphi_{4\,ESD}$	1%	-0.007	0.056**	0.000
	2%	-0.007	0.017	0.003
	5%	0.000	-0.002	0.001
	10%	0.000	0.002	0.001
	90%	0.009*	0.000	0.002
	95%	0.020**	-0.002	0.006
	98%	0.046***	-0.005	0.010**
	99%	0.050***	-0.005	0.031***

Table 2.6 reports tests for contagion from US stock market returns to Colombian pension fund returns (*CPF*), Latin American stock index returns (*LAC*) and the USD/Colombian peso exchange rate (*COPFX*) over three crisis phases by quantile. Coefficient estimates are obtained from regressing the proportion of conditional variance of returns to each local index due to US stock market shocks on a constant and indicators for crisis phases, $F^{-1}(\tau|u_{sys,i,t}) = \varphi_0(\tau) + \varphi_1(\tau)F^{-1}(\tau|u_{sys,i,t-1}) + \varphi_2(\tau)I_{subp,t} + \varphi_3(\tau)I_{CCC,t} + \varphi_4(\tau)I_{ESD,t} + e_{i,t}$ where τ indicates the quantile. Significant coefficients indicate contagion at the *90%, **95%, ***99% confidence level.

2.9 Conclusion

Here we study the behaviour of a particularly isolated (autarchic) asset during the recent financial and sovereign debt crises. Colombian private pension funds can be seen as autarchic assets due to the strict regulatory constraints on their portfolio holdings that confine them largely to defensive, domestic currency assets. Even though these "quantitative restrictions" can limit the risk/return potential of the autarchic portfolio, these same restrictions could also limit the potential losses in times of crisis. Preliminary analysis shows that the Sharpe ratios of private pension fund returns were higher than those of regional and global benchmarks during the crisis period.

We dig deeper into this question by estimating an DCC-GJRGARCH structure with US stock market shocks as the systematic factor and potential source of contagion. We decompose risk into its systematic and idiosyncratic components and test for additional contagious linkages to pension-fund-returns volatility. We also introduce quantile autoregression to overcome some of the limitations and biases generated by asymmetric data and to obtain a more detailed analysis of systemic risk.

Although the coefficient on the US equity risk factor increased in each of the crisis periods, we find no evidence of volatility contagion to Colombian pension funds (CPF) from US equity shocks during the first two phases of the subprime crisis. However, during the European sovereign debt crisis, there is strong evidence of contagion. When we allow for different channels of transmission such as the EMBI and the EFFA fixed-interest indices, there is evidence of volatility contagion from fixed-interest markets to pension fund volatility in the subprime and ESD crises. Therefore, contagion effects seem to be stronger when coming from developed markets than emerging ones, as expected from this type of crisis, where the source is well identified. We also demonstrate contagion to LAC regional stocks during the subprime and CCC episodes.

Our findings are similar to Dooley and Hutchison (2009) in confirming evidence of temporary decoupling during 2007. Our results are also in line with the finding of Boyer et al. (2006) that interdependence among accessible assets was greater than the inaccessible (autarchic) assets. Although there was evidence of decoupling of the Colombian funds in the first and second crisis episodes, our findings also show a strong evidence of recoupling at a later stage of the crisis, in line with those of Frank and Hesse (2009) in their study of the EMBI.

In other words, while restricting pension fund asset holdings to domestic-currency defensive assets may quarantine returns from extreme overseas stock market turbulence, we find no evidence that turmoil in even relatively unrelated fixed-interest markets can be kept at bay. Finally, and most importantly, we hope that by analysing the effect of government regulation in emerging markets we can shed some light on whether and to what extent the effects of contagion can be mitigated.

Chapter 3 The performance of state-owned enterprises in BRIC countries during the GFC

3.1 Introduction

In recent years, emerging economies have moved from being only recipients of foreign direct investment (FDI) to being important investors in their own right. UNCTAD (2012) estimated that in 2010, 30 per cent of global FDI came from emerging economies. At the same time, state-owned enterprises (SOEs) have become significant vehicles for FDI, and, consequently, agents for fostering economic growth and for developing functional capital markets in their countries of origin. Although state-owned multinationals represent just 1 per cent of the total market capitalisation of multinational companies around the world, they account for at least 11 per cent of total FDI. This substantial share is partly due to governments actively and directly promoting economic growth. With an estimated US$5 trillion of assets under management, the share of SOEs in global FDI is likely to continue to increase (UNCTAD, 2012).

We contribute to the literature on the role and performance of SOEs by finding a meaningful mechanism to benchmark SOE performance and test for a "cushion" effect associated with government ownership. In addition, we use quantile regression to analyse the full distribution of factor loadings on SOE returns, including extreme conditions. The empirical analysis based on 70 state-owned enterprises from the BRIC countries and 441 comparable firms from the US demonstrates that government ownership indeed provides protection and thus a "cushion effect" for companies in specific sectors.

The role of SOEs in economic and capital market development raises crucial questions for both academics and policy makers. What, if any, are the effects of government ownership on the financial performance of SOEs? How are investors affected as SOEs are partially privatised and their stock is traded in capital markets around the world?

Most of the literature evaluating SOEs compares the performance of government versus private ownership. Hart et al. (1997) and Shleifer (1998) argue that, from a contractual perspective, a government should be indifferent between either class of ownership since the government can use its regulatory power to draft optimal contracts and force a private supplier to deliver exactly what it wants. Private ownership should therefore be more efficient than government ownership because the private agent will maximise profits subject to the optimal contract. On the other hand, earlier theoretical work by Shapiro and Willig (1990) argued that some products or services have a non-contractible quality that favours public ownership, since the cost of drafting, implementing and monitoring effective regulation that maximises social welfare will outweigh the benefits of privatisation. Finally, Sappington and Stiglitz (1987) argue that, depending on the kind of service or good required by the government, the choice of public versus private ownership is not dichotomous, and that in most cases a public-private partnership should be the optimal solution for the cost-regulation dilemma.

Motivated by this theoretical tension, many empirical studies have tested the comparative performance of state and private ownership, often using data from the privatisation of SOEs (Bortolotti and Perotti, 2007). The first empirical studies on the subject tested relative performance by comparing a set of accounting variables of privatised SOEs against the same variables from their pre-privatisation period. They found that, for both developed and developing countries, privatisation had led to increases in sales, operating efficiency, and profitability (see: Boubakri and Cosset, 1998; D'Souza and Megginson, 1999; Megginson et al., 1994). However, using a comprehensive dataset of Mexican SOEs, La Porta and López-de-Silanes (1999) found that the superior financial performance of SOEs after privatisation was mainly explained by savings from layoffs. Other studies confirm higher profitability and improved total factor productivity (Dewenter and Malatesta, 2001; Estrin et al., 2009). Privatisation of SOEs also helps develop capital markets in less developed economies by attracting new investors (Subrahmanyam and Titman, 1999) and by enabling share issues (Megginson et al., 2004).

Underperformance of government-owned enterprises may be the result of patronage (Shleifer, 1998; Shleifer and Vishny, 1994). Patronage occurs when a government uses its dominant position in SOEs to provide benefits to political supporters, such as above-market wages or bribes and favouritism (La Porta et al., 2002; Dinc, 2005; Nguyen and van Dijk, 2012). Furthermore, politically connected firms are more likely to be bailed out during crises and have more access to external financing than non-connected firms (Faccio et al., 2006) or may enjoy access to lower cost of finance because of implicit government guarantees (Knyazeva *et al.*, 2009; Borisova and Megginson, 2011).[17] On the other hand, there is evidence that partial privatisation can enhance operating performance because of demands for credible financial information from minority shareholders that trade in the public stock market (E. F. Fama, 1980; Holmstrom and Tirole, 1993). Similarly, Chen et al.

[17] Similar evidence of patronage is presented by Dinc and Gupta (2011) for India and Firth et al. (2010) for China. For an alternative perspective on China, see McGuinness (2012), who argued that in the case of the Chinese SOE initial public offerings (IPOs) there was no convincing evidence of underpricing due to an investor perception of higher risk or that politically connected "cornerstone" investors made superior returns during these IPOs.

(2009) argue that in China, the partially privatised SOEs where the central government is the major shareholder exhibited better operating performance than other forms of ownership such as state management bureaus, local governments and private companies, possibly because they provide better safeguards for shareholder wealth protection than their private counterparts (Li et al., 2011).[18] There is evidence for and against public ownership of enterprises: the benefits and costs are likely to be conditional on the business itself, the transparency of decision making and reporting, and the economic and financial context.

Unlike previous work on the subject of privatization, this chapter focuses on the comparative return performance of the largest SOEs in Brazil, Russia, India and China which are commonly known as the BRIC countries against industry competitors in the US during the GFC and what we can learn from the performance of SOEs stocks during the GFC. In this chapter as opposed to the previous one, our main objective is to test for evidence of financial contagion in the context of SOE portfolio performance.. In this chapter as opposed to the previous one, our main objective is to test for evidence of financial contagion in the context of SOE portfolio performance. Here we drill down into the comparative performance of SOEs by allowing for differences during stable and turbulent financial conditions. If government ownership is indeed beneficial in partially privatised companies, providing a form of implicit guarantee, this should be evident in times of crisis, so that SOEs should experience less severe losses than private competitors. In other words, we expect government ownership to act like a "cushion" during crisis

[18] However, Hossain et al. (2013) show that government ownership of financial firms in the Asian region reduces losses in crisis periods but also hinders growth in normal periods, and Liu and Siu (2011) provided evidence that SOEs valued capital investments at a much lower discount rate than partially privatised SOEs and private firms in order to foster growth according to the central government directives.

periods. On the other hand, SOEs will exhibit lower returns than their private counterparts in growth periods or booms because bankruptcy risks are mitigated by stable dividend policies and a more conservative approach toward investment opportunities than private companies (He et al., 2012).

Unlike previous work, we look into the comparative return performance of the largest SOEs in Brazil, Russia, India, and China (BRIC countries) against industry peers in the US. We model the returns of the SOEs and comparable US benchmarks with the standard four-factor model (Carhart, 1997; Eugene F. Fama and French, 1993, 1998). The approach treats US factors as common global factors for the SOE, thus avoiding possible endogeneity issues. And since the US was the originator of the subprime crisis and its aftermath, we hypothesise that US factors will account for contagion to emerging markets.[19]

The remainder of this paper is structured as follows. In Section 3.2 we describe the data and in Section 3.3 we detail the performance models. In Section 3.4 we discuss the results obtained from the linear, quantile, and cross-sectional regression models, and Section 3.5 concludes.

3.2 Data

In order to compile our publicly traded state-owned enterprises dataset and avoid liquidity issues, we selected only SOEs that are components of their national stock market indexes[20] and where the majority of the company shares are controlled by the

[19] For examples of recent studies of the US as carrier of contagion see: (Beirne et al., 2008; Bekaert, Ehrmann, et al., 2011; Bekaert, Harvey, et al., 2011; Dooley and Hutchison, 2009; Dufrénot et al., 2011; Dungey, Milunovich, et al., 2010b; Fazio, 2007; Felices and Wieladek, 2012; Korkmaz et al., 2012).

[20] In the case of Brazil we used the BOVESPA, India the BSE 30 index, and Russia the RTS Index. In the case of China we used the CSISOE which is a special index that comprises the 40 largest state-owned enterprises in mainland China.

central government or any other form of government agency.[21] For all the companies in our study, with the exception of the consumer staples sector in Brazil, the state holds majority control (50 per cent or more). In the consumer staple sector in Brazil, the state holdings add up to only 40 per cent, but these are still larger than those held by private shareholders, so the state can exert considerable control in the companies. To identify US private counterparts to the SOEs, we selected stocks that comprised the matching S&P 500 Global Industry Classification Standard (GICS) level 1 indexes. SOE and index data are daily closing prices converted to USD from Bloomberg for the period January 3, 2000, to April 30, 2012. The Fama and French (FF) factors plus momentum were downloaded from Kenneth French's website. We use short-term (weekly) and medium-term (monthly) holding periods to compute portfolio returns, and weekly and monthly log returns for each stock. Therefore, the equally weighted portfolio returns series include 649 weekly and 148 monthly observations respectively. The only exception is the consumer discretionary sector that includes one Chinese SOE that began to publicly trade in the mid-2000s, with 337 weekly and 77 monthly observations. Table 3.1 reports summary statistics for the SOEs' stocks and benchmarks by industry sector.

[21] For example state asset management bureaus or development banks.

Table 3.1: Number of SOEs and comparables by GICS industry sector

		State Owned Enterprises					
GICS	**Sector**	**Brazil**	**Russia**	**India**	**China**	**Total**	**Comparables**
25	Consumer discretionary				1	1	80
30	Consumer staples	3			3	6	39
10	Energy	1	1	2	7	11	43
40	Financials	1	1	3	15	20	80
20	Industrials	2	1	1	6	10	61
45	Information technology				1	1	69
15	Materials	1			4	5	31
50	Telecommunication services	1			3	4	8
55	Utilities	4	4	3	1	12	30
	Total	**13**	**7**	**9**	**41**	**70**	**441**

3.3 Model

The key performance variable aims to capture the component of the return that is due to the stake of the government ("state"). If the difference between the SOE return and a comparable firm that is not state-owned is significantly positive, the SOE outperforms the comparable firms and *vice versa*. We create a performance variable for each industry sector as $SECSOE_t$-$COMPSEC_t$ where $SECSOE_t$ is the return to an equally weighted portfolio of the SOE's stocks in each sector net of the risk-free rate, minus the return to an equally weighted portfolio of US comparables ($COMPSEC_t$) in the same sector, net of the risk-free rate.[22] (Our method follows Hong and Kacperczyk's (2009) valuation of sin stocks.)

The control variables are the FF factors[23] and momentum: 1) $MKTPREM_{t,}$ which is the excess market return of a portfolio of firms listed in the NYSE, AMEX, and NASDAQ stock exchanges over the risk free rate; 2) SMB_t, (small minus big), which measures the size premium so that a positive coefficient means that small

[22] We use the one-month US Treasury bill as a proxy for this rate.
[23] For detailed information on how to construct the factors please refer to Kenneth R. French's website.

capitalisation stocks outperform large capitalisation and *vice versa*; 3) HML_t (high minus low), which measures the premium to investing in companies with a high book-to-market ratio, so that a positive coefficient means that “value” outperforms “growth” companies in a given observation period; 4) MOM_t or momentum, which is constructed by building portfolios that are long on stocks with the highest returns and short on stocks with the lowest returns during the last two to 12 months, so that a positive coefficient means that past winners outperformed past losers and *vice versa* (Daniel et al., 2012; Gutierrez and Gaglianone, 2008).

Also, in order to observe the effect of changing regimes during the financial crisis that started in 2007, we divided the crisis into three phases denoted by indicator variables. The first-phase "subprime" crisis (D_{Sub}) begins July 26, 2007, which was the day the Dow Jones recorded a large loss in response to bad news from mortgage lender Countrywide Financial. At this point, the market processed news of “difficult conditions” in the subprime market following Countrywide Financial Corporation’s SEC filing on July 24. The beginning of the "credit crunch" crisis (D_{Lehman}) is generally dated from when Lehman Brothers filed for bankruptcy on September 15, 2008. We date the European sovereign debt crisis (D_{ESD}) from October 22, 2009, when Fitch first downgraded and reported a negative outlook for Greek sovereign debt. The turbulence in European debt markets was continuing at the end of our sample, April 30, 2012.[24]

The conditional mean equation for the four-factor model is:

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[24] The key dates for the subprime and credit crunch crises were taken from the financial turmoil timeline chart from the Federal Reserve Bank of St Louis (http://timeline.stlouisfed.org/pdf/CrisisTimeline.pdf) and for the European sovereign debt crisis from the credit rating function in Bloomberg. There are other studies that use similar dates for the CCC and place the subprime around the same period, including Frank and Hesse (2009), Dooley and Hutchison (2009), and Felices and Wieladek. (2012).

$$(SECSOE_t - COMPSEC_t) = \alpha_t + \beta_1 MKTPREM_t + \beta_2 SMB_t + \beta_3 HML_{i,t} + \beta_4 MOM_t + \gamma_1 D_{Sub} + \gamma_2 D_{Lehman} + \gamma_3 D_{ESD} + \varepsilon_t \quad (3.1)$$

If the SOE firms perform better (worse) than their benchmarks after controlling for the four factors and the crisis indicators, the coefficient α will be positive (negative). However, we are particularly interested in the performance of SOEs during crisis periods. If SOEs outperform their benchmarks during a financial crisis the government's stake in the firm is likely to be cushioning the value of the firm. The economic rationale for this "cushion" effect is that the stake of the government is either so large that the firm is fully protected from general market conditions or that investors in state-owned firms know that the government provides a certain degree of protection and thus are more reluctant to sell the shares despite the crisis conditions.

If SOEs outperform their benchmarks during the entire crisis period, all three coefficients on crisis indicators in equation (3.1) will be positive. If the outperformance, and thus the cushion effect, is only evident in one of the three crisis periods, it will show in a positive gamma coefficient representing that specific crisis period. Since we *do not* test whether the correlation between the SOEs and the benchmarks or the market has changed, as is a premise in the contagion literature, the model given by equation (3.1) is not a test of contagion. It is primarily a test of a cushion effect. Additionally, by analysing the performance of the dependent variable for different quantiles, we can obtain a more detailed picture of the relationships. We use the quantile regression model as proposed by (Koenker and Bassett Jr, 1978):

$$\begin{aligned} F^{-1}(\tau | SECSOE_t - COMPSEC_t) &= \alpha_t(\tau) + \beta_1(\tau) MKTPREM_t + \beta_2(\tau) SMB_t + \beta_3(\tau) HML_{i,t} \\ &+ \beta_4(\tau) MOM_t + \gamma_1(\tau) D_{Sub} + \gamma_2(\tau) D_{Lehman} + \gamma_3(\tau) D_{ESD} + F_\varepsilon^{-1}(\tau | SECSOE_t - COMPSEC_t) \end{aligned} \quad (3.2)$$

where τ represents the 1%, 5%, 50%, 95%, and 99% quantiles for the weekly[25] and 3%, 5%, 50%, 95%, and 97% quantiles for the monthly observations. By estimating over the extreme quantiles we can observe any asymmetric structure of the cushion effect in the tails of the distribution. By analysing the gamma coefficients across quantiles we can test if indeed certain SOE industry sectors exhibit a "cushion effect" not only on average, i.e., in the mean, but also in the tails of the distribution.

We run a cross-sectional regression model to assess the performance of SOEs against comparables, and account for fixed effects using a similar model to that proposed by Hong and Kacperczyk (2009). We regress the reported cross-sectional average excess return to each stock during the observation period on its own lag, the lagged average change in market-to-book[26] value (AVGBK) and the lagged average change in market capitalisation (AVGCAP). The conditional excess return specification is:

$$AVGER_{it} = c_0 + c_1 AVGER_{it-1} + c_2 AVGCAP_{it-1} + c_3 AVGBK_{it-1} + \mathbf{c_4 D_{it}} + \varepsilon_{it},\ i = 1,....n \tag{3.3}$$

In this case the coefficient of interest is the vector of loadings on the dummy variables (D_{it}) that account for all possible combinations of fixed effects for industry classifications as well as the SOE's country of origin. The next section describes the results.

[25] In the case of the energy and discretionary SOE sectors the weekly and monthly quantiles are: 3%, 5%, 10%, 50%, 95%, and 97% and 5%, 8%, 25%, 50%, 75%, and 92% respectively , since there are not enough data points to compute higher quantiles**.**

[26] In this specific case we use the inverse of the book-to-market ratio as reported by Bloomberg.

3.4 Results

3.4.1 Results of the four-factor model linear regression

The results of the four-factor linear regressions for all the sectors are summarised in Table 3.2. The MKTPREM based on the US market index is significant in all but the industrial and material sectors, but, with the exception of the financial sector, there are no significant alphas (α).

Table 3.2: Results of the four-factor linear regression model

The results are the estimated coefficients obtained from the following OLS regression:

$$(SECSOE_t - COMPSEC_t) = \alpha_t + \beta_1 MKTPREM_t + \beta_2 SMB_t + \beta_3 HML_{i,t} + \beta_4 MOM_t + \gamma_1 D_{Sub} + \gamma_2 D_{Lehman} + \gamma_3 D_{ESD} + \varepsilon_t$$

Where $(SECSOE_t - COMPSEC_t) =$ an equally weighted portfolio of the corresponding SOE sector stocks that compose the sector net of the risk-free interest rate minus an equally weighted portfolio of US-sector comparable stocks net of the risk-free interest rate for the period between 2000 and 2012 on a weekly and monthly basis. MKTPREM, SMB, HML, and MOM are the US explanatory factors downloaded on a weekly and monthly basis from the Kenneth R. French website. D_{Sub}, $L_{ehman,}$ and D_{ESD} are dummy variables that take the value of 0 or 1 for our specified crisis periods as defined in Section 3.3. All standard errors are adjusted using the Newey-West correction for serial correlation. ***1%; **5%; and *10% significance.

WEEKLY (2000–12)						MONTHLY (2000–12)					
SOE	**ALPHA**	**MKTPREM**	**SMB**	**HML**	**MOM**	**SOE**	**ALPHA**	**MKTPREM**	**SMB**	**HML**	**MOM**
DISCRETIONARY	0.0089	0.2998	-0.5350	-0.2779	0.3320	DISCRETIONARY	0.0328	1.0507	-1.2557	-1.9255	-0.4902
p-value	(0.1905)	(0.1120)	**(0.0997)** *	(0.4389)	**(0.0767)** *	p-value	(0.4813)	**(0.0013)** ***	(0.1052)	**(0.0005)** ***	**(0.0253)** **
STAPLES	0.0022	0.4112	0.5309	0.0679	-0.0271	STAPLES	0.0163	1.2321	0.4282	-0.8414	0.4567
p-value	(0.5071)	**(0.0002)** ***	**(0.0044)** ***	(0.5864)	(0.8001)	p-value	(0.1979)	**(0.0000)** ***	(0.3421)	**(0.0500)** **	(0.1186)
ENERGY	-0.0004	-0.3240	0.1717	-0.2753	-0.2497	ENERGY	-0.0029	0.0342	0.0790	-0.2796	0.0048
p-value	(0.8614)	**(0.0000)** ***	(0.2957)	**(0.0286)** **	**(0.0036)** ***	p-value	(0.7884)	(0.8536)	(0.7993)	(0.2128)	(0.9747)
FINANCIAL	0.0032	-0.3180	0.2728	-0.6574	0.1711	FINANCIAL	0.0169	0.3682	0.0837	-0.9923	0.2847
p-value	(0.1183)	**(0.0013)** ***	**(0.0542)**	**(0.0000)** ***	**(0.0776)** *	p-value	**(0.0245)** **	**(0.0342)** **	(0.7258)	**(0.0008)** ***	**(0.0315)** **
INDUSTRIAL	0.0001	-0.1103	0.1788	-0.0883	-0.0169	INDUSTRIAL	0.0039	0.1852	0.0807	-0.4821	0.1313
p-value	(0.9313)	(0.1512)	(0.1389)	(0.2861)	(0.8021)	p-value	(0.5367)	(0.1878)	(0.6187)	**(0.0020)** ***	(0.2540)
INFORMATION	-0.0062	-0.3916	0.1123	0.7851	0.1299	INFORMATION	-0.0182	-0.2437	0.1890	0.2979	0.2374
p-value	(0.1113)	**(0.0105)** **	(0.6575)	**(0.0009)** ***	(0.3488)	p-value	(0.3720)	(0.6071)	(0.7806)	(0.6032)	(0.5257)
MATERIALS	0.0005	-0.0282	0.6662	-0.0415	0.0315	MATERIALS	0.0082	0.6843	0.2879	-0.7109	0.2244
p-value	(0.9015)	(0.8175)	**(0.0011)** ***	(0.8449)	(0.7597)	p-value	(0.6693)	**(0.0091)** *	(0.2611)	**(0.0169)** **	(0.2017)
TELECOMMUNICATIONS	0.0012	-0.0393	0.2829	-0.3475	0.2362	TELECOMMUNICATIONS	0.0094	0.4450	0.1365	-0.6594	0.1921
p-value	(0.6449)	(0.6562)	**(0.0978)** *	**(0.0374)** **	**(0.0112)** **	p-value	(0.3438)	**(0.0571)** *	(0.5788)	**(0.0220)** **	(0.2799)
UTILITIES	0.0005	0.2667	0.7164	-0.2648	-0.0798	UTILITIES	0.0068	0.8678	0.3606	-0.5999	-0.0066
p-value	(0.8492)	**(0.0157)** **	**(0.0001)** ***	(0.1596)	(0.5276)	p-value	(0.4329)	**(0.0007)** ***	(0.2028)	**(0.0037)** ***	(0.9712)

In the case of the discretionary sector in the short term (weekly data) the most significant factors are the SMB and MOM, whereas in the medium term (monthly data) the SOE discretionary sector excess returns can be explained by comparable US growth (low book to market ratio) stocks and a contrarian strategy (negative momentum). The consumer staples sector shows a similar pattern.

The energy SOE sector shows a countercyclical pattern where weekly returns in this sector are inversely related to market returns, and explained by the US growth factor and a contrarian strategy. The financial SOE sector is of special interest during this sample period when US financial-sector influences were dominant around the world: the MKTPREM exhibits a negative coefficient and all four factors are significant using weekly data. For monthly data, the financial SOE sector exhibits a significantly positive α and three of the four factors are significant.

Sectorial differences are also noticeable for the remaining categories, and a few are independent of the equivalent US factors. Materials, telecommunications, and utilities SOE returns are explained by US benchmarks, but industrial and information technology SOE returns are mainly independent of US factors.

The extent of shifts in alpha during crisis periods is reported in Table 3.3. The pattern of results reflects the sources of uncertainty during different phases of the crisis. None of the SOEs' alphas are significantly affected by the first-phase subprime crisis when the trouble appeared to be contained to the financial and real estate sectors. During the second-phase credit crunch, the financial sector SOEs show a superior performance relative to the US comparables, most probably because emerging economies were less affected by the recession. Finally, during the third-phase European sovereign debt episode, when sovereign risk was most prominent, coefficients on the indicator variables for all sector SOEs, with the exception of information technology, are negative. In the financial and consumer staples sectors the effect is significant for both weekly and monthly returns. For the industrials, materials, and utilities sectors the effect is more relevant in the monthly models whereas the telecommunication services sector seems immune to the crisis phases, at least at the conditional mean return level.

Overall, we observe that SOEs cannot be lumped together – factor sensitivities vary in sign and size – and while there is no evidence for average alpha over the sample, SOE conditional expected excess returns will deviate both above and below zero during crises.

Table 3.3: Crisis effect by sector

The results are the estimated coefficients obtained from the following OLS regression:

$$(SECSOE_t - COMPSEC_t) = \alpha_t + \beta_1 MKTPREM_t + \beta_2 SMB_t + \beta_3 HML_{t,t} + \beta_4 MOM_t + \gamma_1 D_{Sub} + \gamma_2 D_{Lehman} + \gamma_3 D_{ESD} + \varepsilon_t$$

Where $(SECSOE_t - COMPSEC_t)$ = an equally weighted portfolio of the corresponding SOE sector stocks that compose the sector net of the risk-free interest rate minus an equally weighted portfolio of US-sector comparable stocks net of the risk-free interest rate for the period between 2000 and 2012 on a weekly and monthly basis. MKTPREM, SMB, HML, and MOM are the US explanatory factors downloaded on a weekly and monthly basis from the Kenneth R. French website. D_{Sub}, $L_{ehman,}$ and D_{ESD} are dummy variables that take the value of 0 or 1 for our specified crisis periods as defined in Section 3.3. All standard errors are adjusted using the Newey-West correction for serial correlation. ***1%; **5%; and *10% significance

	Discretionary		**Staples**		**Energy**		**Financials**	
	Weekly	Monthly	Weekly	Monthly	Weekly	Monthly	Weekly	Monthly
(D_{sub})	-0.0129	-0.0099	-0.0040	-0.0062	-0.0044	-0.0048	-0.0065	-0.0071
p-value	(0.2616)	(0.8579)	(0.5540)	(0.7962)	(0.5158)	(0.7486)	(0.2939)	(0.6705)
(L_{ehman})	0.0152	0.0284	0.0035	0.0154	0.0047	0.0238	0.0110	0.0319
p-value	(0.1379)	(0.6098)	(0.6120)	(0.6219)	(0.4348)	(0.3136)	**(0.0184)** **	(0.1688)
(D_{ESD})	-0.0101	-0.0427	-0.0077	-0.0520	-0.0008	-0.0100	-0.0069	-0.0371
p-value	(0.2154)	(0.4016)	**(0.0817)** *	**(0.0083)** ***	(0.8127)	(0.7449)	**(0.0351)** **	**(0.0016)** ***

	Industrials		**Infotec**		**Materials**		**Telco**	
	Weekly	Monthly	Weekly	Monthly	Weekly	Monthly	Weekly	Monthly
(D_{sub})	-0.0056	-0.0175	0.0054	0.0259	-0.0090	-0.0192	0.0019	0.0162
p-value	(0.3198)	(0.4309)	(0.5839)	(0.3430)	(0.3359)	(0.5998)	(0.7835)	(0.5929)
(L_{ehman})	0.0037	0.0117	0.0096	0.0158	0.0080	0.0221	-0.0027	0.0014
p-value	(0.4617)	(0.6548)	(0.4197)	(0.8571)	(0.2072)	(0.3762)	(0.4752)	(0.9581)
(D_{ESD})	-0.0048	-0.0280	0.0093	0.0050	-0.0076	-0.0474	0.0053	-0.0161
p-value	(0.1218)	**(0.0224)** **	**(0.0928)** *	(0.9118)	(0.0833)	**(0.0300)** **	(0.4540)	(0.2931)

	Utilities	
	Weekly	Monthly
(D_{sub})	0.0011	0.0180
p-value	(0.8711)	(0.1725)
(L_{ehman})	0.0034	0.0008
p-value	(0.6885)	(0.9811)
(D_{ESD})	-0.0045	-0.0297
p-value	(0.1595)	**(0.0154)** **

3.4.2 Results of the four-factor model quantile regressions

Quantile regression estimates the explanatory power of the factors for extremely high and low returns, as well as estimating the shifts in mean returns during the phases of the crisis. Table 3.4 reports quantile regression results for the coefficient on the market factor for each SOE sector. For weekly returns, the coefficient on the market factor is significant at different quantiles in the majority of the sectors with the exception of the telecommunication sector, whereas for monthly data, the coefficient is insignificant for industrials, information technology, and energy sectors. For the financials, industrial, and energy sectors the sign of the coefficients is negative. The market premium is significant for extreme negative and positive quantiles in the consumer staples, financials, materials, energy, and industrial (large positive returns only) sectors for the weekly returns and extreme negative returns in the consumer staples, utilities, and telecommunication sector and extreme positive excess returns in the staples, materials, and consumer discretionary sectors.

Table 3.4: Quantile regression estimates for weekly and monthly returns – CAPM (MKTPREM)

By generalising equation (3.1) in Section 3.3 we obtain:

$F^{-1}(\tau|SECSOE_t - COMPSEC_t) = \alpha_t(\tau) + \beta_1(\tau)MKTPREM_t + \beta_2(\tau)SMB_t + \beta_3(\tau)HML_{i,t} + \beta_4(\tau)MOM_t + \gamma_1(\tau)D_{Sub} + \gamma_2(\tau)D_{Lehman} + \gamma_3(\tau)D_{ESD} + F_\varepsilon^{-1}(\tau|SECSOE_t - COMPSEC_t)$ where (τ) = the 1%, 5%, 50%, 95%, and 99% for the weekly and 3%, 5%, 50%, 95%, and 97% for the monthly observations.

MKTPREM COEFFICIENTS-WEEKLY (2000–2012)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
1%	0.3842**	-0.2040	0.3288	-0.4539	1.0903***	0.0935	0.0973	0.3848	-0.3739*
5%	0.5589***	-0.3614**	-0.0681	0.0017	1.1395***	0.0024	0.1083	0.3558	-0.3853*
50%	0.3860***	-0.3239***	-0.1947***	-0.3048***	1.1538***	-0.0918	0.2700***	0.3235*	-0.3094***
95%	0.4231***	-0.4021***	-0.2956**	-0.5177	0.9957***	-0.0432	0.2538	0.2190	-0.3644***
99%	0.4842**	-0.5812***	-0.0145	-1.1036	0.8251***	0.0412	0.0587	0.0700	-0.3046**

MKTPREM COEFFICIENTS-MONTHLY (2000–2012)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
3%	1.2418***	-0.1728	-0.2850	-1.6818	0.5621	0.7264	0.8245**	0.1895	0.2215
5%	1.2533***	0.2060	-0.1386	-0.2751	0.3201	1.1180**	1.0071**	0.2912	0.3644
50%	0.7502**	0.4848**	0.1380	-0.2437	0.4522*	0.2237	1.4116***	1.0614**	-0.0177
95%	1.0577	0.2879	-0.2032	0.3256	1.0531	0.4977	0.4082	1.6849***	-0.0788
97%	1.4870**	0.2802	-0.2239	0.7541	2.0795**	0.9587	0.5191	1.3834***	-0.0592

+ In the case of the energy and discretionary SOEs sectors the weekly and monthly quantiles should be interpreted as: 3%, 5%, 50%, 95%, and 97% and 5%, 8%, 50%, 75%, and 92% respectively, since there are not enough data points to compute higher quantiles in these two sectors. ***1%; **5%; and *10% significance.

The SMB factor is significant in explaining part of the variation in large negative and positive excess returns in the consumer staples, financials, utilities, and telecommunication services (large negative returns only) SOEs sectors at the weekly level (Table 3.5). At the monthly level, the SMB factor is significant in explaining large negative excess returns in the information technology and materials SOEs sector and positive excess returns in the consumer discretionary sector.

Table 3.5: Quantile regression estimates for weekly and monthly holding periods – small minus big factor (SMB)

By generalising equation (3.1) in Section 3.3 we obtain:
$F^{-1}(\tau|SECSOE_t - COMPSEC_t) = \alpha_t(\tau) + \beta_1(\tau)MKTPREM_t + \beta_2(\tau)SMB_t + \beta_3(\tau)HML_{i,t} + \beta_4(\tau)MOM_t + \gamma_1(\tau)D_{Sub} + \gamma_2(\tau)D_{Lehman} + \gamma_3(\tau)D_{ESD} + F_{\varepsilon}^{-1}(\tau|SECSOE_t - COMPSEC_t)$ where (τ) = the 1%, 5%, 50%, 95%, and 99% for the weekly and 3%, 5%, 50%, 95%, and 97% for the monthly observations.

SMB COEFFICIENTS: WEEKLY (2000–12)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
1%	1.6089***	1.2574***	0.5450	0.0057	0.0007	0.7876*	1.6096***	-0.0808	0.3769
5%	1.1619***	0.5371**	0.3974	-0.1707	0.0696	-0.0007	1.0673***	-0.3182	0.5171
50%	0.2351	0.0641	0.1114	-0.1012	0.0937	0.1296	0.5760***	-0.2915	0.1411
95%	0.5052	0.3297	0.2221	1.1065	-0.0302	0.1596	0.8387**	-0.7580	0.1429
99%	1.1033*	0.8056**	0.0911	1.6645	0.0093	-0.0300	0.3761	-0.6229	-0.0465

SMB COEFFICIENTS: MONTHLY (2000–12)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
3%	-0.4947	-0.0670	0.9716	4.8511*	1.3094	0.0091	0.5647	-0.8626	0.2729
5%	-0.3941	0.1450	0.2405	1.4180	1.2331*	-0.1752	0.1654	-0.2841	0.3268
50%	0.1710	0.2182	-0.0469	0.1890	0.2087	-0.1152	0.1367	-1.7586*	-0.2555
95%	0.7513	0.0656	-0.0600	0.3418	-0.6069	0.3376	0.1793	-0.3579	0.4425
97%	1.3314	0.0271	-0.1153	0.1562	-0.4700	0.5136	0.5729	-2.4097**	0.4306

+ In the case of the energy and discretionary SOEs sectors the weekly and monthly quantiles should be interpreted as: 3%, 5%, 50%, 95%, and 97% and 5%, 8%, 50%, 75%, and 92% respectively, since there are not enough data points to compute higher quantiles in these two sectors. ***1%; **5%; and *10% significance.

From Table 3.6 we can observe that for both the weekly and monthly returns, the HML factor coefficients are significant in most of the SOEs sectors with the exception of consumer staples and energy. For financial sector SOEs, the HML factor coefficients have negative signs, implying a similarity with US growth stocks. The HML factor is significant in explaining part of the variation in large negative and positive excess returns in the financials, information technology, and materials at the weekly level. In the case of telecommunication services, utilities, and consumer discretionary the factor explains part of the variation associated with large positive excess returns. At the monthly level, the HML factor is significant in explaining large negative excess returns in the information technology and materials SOEs sector and positive excess returns in the consumer discretionary, industrials, and telecommunication services sector.

Table 3.6: Quantile regression estimates for weekly and monthly holding periods – high minus low factor (HML)

By generalising equation (3.1) in Section 3.3 we obtain:
$F^{-1}(\tau | SECSOE_t - COMPSEC_t) = \alpha_t(\tau) + \beta_1(\tau) MKTPREM_t + \beta_2(\tau) SMB_t + \beta_3(\tau) HML_{i,t} + \beta_4(\tau) MOM_t + \gamma_1(\tau) D_{Sub} + \gamma_2(\tau) D_{Lehman} + \gamma_3(\tau) D_{ESD} + F_{\varepsilon}^{-1}(\tau | SECSOE_t - COMPSEC_t)$ where (τ) = the 1%, 5%, 50%, 95%, and 99% for the weekly and 3%, 5%, 50%, 95%, and 97% for the monthly observations.

HML COEFFICIENTS: WEEKLY (2000–12)									
Quantile	**Staples**	**Financials**	**Industrials**	**Infotec**	**Materials**	**Telco**	**Utilities**	**Discretionary+**	**Energy+**
1%	-0.0333	-0.4828	0.6410	2.1905***	0.5025*	0.2868	0.2544	-0.1905	-0.3823
5%	-0.1673	-0.7550***	0.0975	1.1924*	0.6812***	-0.4006	0.0643	-0.4318	-0.1511
50%	-0.0345	-0.6112***	-0.1286	0.7097***	0.5480***	-0.2897*	-0.3389*	-0.0513	-0.2050
95%	-0.0712	-0.6536**	0.0332	0.4709	0.8249***	-0.6477**	-0.2803	-1.2947*	-0.0880
99%	0.0244	-0.6199**	-0.2162	2.6846**	0.6745***	-0.5134	-0.6619*	-1.7998**	-0.1658

HML COEFFICIENTS: MONTHLY (2000–12)									
Quantile	**Staples**	**Financials**	**Industrials**	**Infotec**	**Materials**	**Telco**	**Utilities**	**Discretionary+**	**Energy+**
3%	-0.7035	-0.8624	0.2850	2.2237**	-0.6470	-0.4980	-0.4964	0.2015	-0.7377
5%	-0.5146	-0.6252	-0.3168	0.4505	-1.0741**	-0.6276	-0.8406	-0.6799	-0.6871
50%	-0.5250	-1.3345***	-0.4439*	0.2979	-0.6533*	-0.4987	-0.3871	-2.1925***	-0.5094
95%	-0.9738	-0.8481	-0.8242**	0.3248	-0.2601	-1.3414**	-0.5512	-2.9834***	-0.0147
97%	-1.8206	-0.6850	-0.8496**	0.4698	-1.0990	-1.4585**	-0.4902	-0.1314	-0.0068

+ In the case of the energy and discretionary SOEs sectors the weekly and monthly quantiles should be interpreted as: 3%, 5%, 50%, 95%, and 97% and 5%, 8%, 50%, 75%, and 92% respectively, since there are not enough data points to compute higher quantiles in these two sectors. ***1%; **5%; and *10% significance.

Estimated coefficients for the momentum factor at different quantile levels are reported in Table 3.7. The MOM factor is significant in explaining part of the variation in large negative and positive excess returns in the telecommunication services sector. In the case of the utilities, consumer discretionary, and energy sectors the factor explains part of the variation associated with large positive excess returns. At the monthly level, the MOM factor is significant in explaining large positive excess returns in the telecommunication services and consumer discretionary sectors. A negative coefficient means that a contrarian strategy in the US benchmark can explain part of the variation in returns. Overall, we see that the four-factor model has explanatory power at all quantiles of excess returns to SOEs, but that the sign and size of the factor coefficients vary both by sector and by quantile.

Table 3.7: Quantile regression estimates for weekly and monthly holding periods – momentum factor (MOM)

By generalising equation (3.1) in Section 3.3 we obtain:

$F^{-1}(\tau|SECSOE_t - COMPSEC_t) = \alpha_t(\tau) + \beta_1(\tau)MKTPREM_t + \beta_2(\tau)SMB_t + \beta_3(\tau)HML_{i,t} + \beta_4(\tau)MOM_t + \gamma_1(\tau)D_{Sub} + \gamma_2(\tau)D_{Lehman} + \gamma_3(\tau)D_{ESD} + F_{\varepsilon}^{-1}(\tau|SECSOE_t - COMPSEC_t)$ where (τ) = the 1%, 5%, 50%, 95%, and 99% for the weekly and 3%, 5%, 50%, 95%, and 97% for the monthly observations.

MOM COEFFICIENTS: WEEKLY (2000–12)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
1%	-0.4395	-0.4153	-0.1102	-0.2752	-0.0906	0.7500***	-0.3644**	0.6486*	-0.7555***
5%	-0.1985	-0.0206	-0.1867*	0.5414	-0.0862	0.3654***	-0.2054	0.5283	-0.5219***
50%	0.0001	0.1561	0.0281	0.2685**	0.1058**	0.2300*	-0.0225	0.2872	-0.1983**
95%	-0.0982	0.0190	0.0769	-0.1120	-0.0173	0.1836	0.0368	-0.2423	-0.2030
99%	0.0608	0.0677	-0.0198	-0.3805	0.0039	0.5655**	-0.1763	-0.1636	-0.1769

MOM COEFFICIENTS: MONTHLY (2000–12)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
3%	0.3077	-0.0137	0.0054	0.4354	0.2899	0.8529*	-0.5203	0.0558	-0.4545
5%	0.1960	0.2374	0.1773	0.1675	-0.0682	0.8170	-0.3435	-0.6999**	-0.3088
50%	0.1462	0.4138**	0.1422	0.2374	0.1980	-0.0654	0.3445	-0.4943	-0.0644
95%	0.8298	0.3535	0.1203	0.2237	-0.6753	0.2863	-0.1593	-0.3838	-0.0706
97%	1.0095	0.3897	0.0989	0.2936	-0.1654	0.2076	0.0205	0.0934	-0.0701

+ In the case of the energy and discretionary SOEs sectors the weekly and monthly quantiles should be interpreted as: 3%, 5%, 50%, 95%, and 97% and 5%, 8%, 50%, 75%, and 92% respectively, since there are not enough data points to compute higher quantiles in these two sectors. ***1%; **5%; and *10% significance.

Turning to crisis effects, from estimates in Table 3.8 we see that the coefficients on the subprime crisis indicators are significant for the majority of SOE sectors with the exception of the materials, information technology, and utilities. At a monthly level the effects are much weaker. Large negative excess returns in the staples, industrials, and discretionary SOEs sectors were amplified in the subprime crisis and both positive and negative extremes were amplified in the financial and energy SOEs sectors. However, in some instances, the signs were reversed, so that the crisis dampened the effects of extreme returns. In the telecommunication services (weekly) and utilities (monthly) the lower quantiles are characterised by positive coefficients. This is evidence of a "cushion effect" against extremes and suggests that these sectors were protected from extreme negative movements during the subprime crisis.

Table 3.8: Quantile regression estimates for weekly and monthly holding periods – subprime crisis

By generalising equation (3.1) in Section 3.3 we obtain: $F^{-1}(\tau | SECSOE_t - COMPSEC_t) = \alpha_t(\tau) + \beta_1(\tau) MKTPREM_t + \beta_2(\tau) SMB_t + \beta_3(\tau) HML_{i,t} + \beta_4(\tau) MOM_t + \gamma_1(\tau) D_{Sub} + \gamma_2(\tau) D_{Lehman} + \gamma_3(\tau) D_{ESD} + F_\varepsilon^{-1}(\tau | SECSOE_t - COMPSEC_t)$ where $(\tau) =$ the 1%, 5%, 50%, 95%, and 99% for the weekly and 3%, 5%, 50%, 95%, and 97% for the monthly observations. In the telecommunication services (weekly) and utilities (monthly) we reject the null of no contagion since the lower quantiles are characterised by positive coefficients. This is evidence of a "cushion effect" against contagion and suggests that these sectors are immune or at least that there is a decoupling in the tails regarding extreme negative movements during the subprime crisis (positive coefficients associated with large negative returns at lower quantiles).

SUBPRIME COEFFICIENTS: WEEKLY (2000–12)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
1%	-0.0282	-0.0613**	-0.0391*	0.0136	-0.0142	0.0375***	0.0023	-0.0164	0.0005
5%	-0.0372**	-0.0260	-0.0429***	-0.0215	-0.0088	0.0232**	0.0153	-0.0356*	-0.0045
50%	0.0012	-0.0068	-0.0087	0.0112	0.0034	-0.0030	0.0034	-0.0116	-0.0076
95%	0.0018	0.0217	0.0280**	0.0224	0.0053	-0.0117	-0.0003	0.0138	0.0086
99%	-0.0101	0.0835**	0.0238	-0.0275	0.0012	-0.0119	-0.0153	0.0296	0.0321*

SUBPRIME COEFFICIENTS: MONTHLY (2000–12)

Quantile	Staples	Financials	Industrials	Infotec	Materials	Telco	Utilities	Discretionary+	Energy+
3%	0.0594	-0.0552*	0.0046	0.1573	0.0704	0.0540	0.1033**	-0.0905	0.0756
5%	0.0331	-0.0267	-0.0308	0.0845	0.0465	0.0299	0.1175***	-0.0105	0.0640
50%	-0.0091	-0.0368	-0.0299	0.0141	-0.0225	-0.0040	-0.0111	-0.0172	-0.0349
95%	-0.0205	0.0290	-0.0067	0.1392	0.0023	-0.0232	0.0170	0.0016	0.0292
97%	-0.1532	0.0176	-0.0095	0.1024	-0.1445	-0.0384	-0.0059	-0.0697	0.0288

+ In the case of the energy and discretionary SOEs sectors the weekly and monthly quantiles should be interpreted as: 3%, 5%, 50%, 95%, and 97% and 5%, 8%, 50%, 75%, and 92% respectively, since there are not enough data points to compute higher quantiles in these two sectors. ***1%; **5%; and *10% significance

In Table 3.9 we can observe the effect of the credit crunch (Lehman) crisis at different quantile levels. For weekly returns, the coefficients of the credit crunch crisis are significant for financials, industrials, materials, and telecommunication services SOEs sectors, whereas for monthly returns, the only significant coefficients are found in the materials SOE sector. In the short term and in the financial and materials SOEs sectors, large positive excess returns and large negative excess returns are both amplified. By contrast, in the telecommunication sector (weekly) and materials (monthly) the lower quantiles are characterised by positive coefficients, and upper quantiles by negative coefficients. This is again evidence of a "cushion effect" during the credit crunch crisis (positive coefficients associated with large negative returns at lower quantiles).

Table 3.9: Quantile regression estimates for weekly and monthly holding periods – credit crunch crisis

By generalising equation (3.1) in Section 3.3 we obtain:
$F^{-1}(\tau | SECSOE_t - COMPSEC_t) = \alpha_t(\tau) + \beta_1(\tau) MKTPREM_t + \beta_2(\tau) SMB_t + \beta_3(\tau) HML_{i,t} + \beta_4(\tau) MOM_t + \gamma_1(\tau) D_{Sub} + \gamma_2(\tau) D_{Lehman} + \gamma_3(\tau) D_{ESD} + F_{\varepsilon}^{-1}(\tau | SECSOE_t - COMPSEC_t)$ where $(\tau) =$ the 1%, 5%, 50%, 95%, and 99% for the weekly and 3%, 5%, 50%, 95%, and 97% for the monthly observations. In the telecommunication services (weekly) and materials (monthly) we reject the null of no contagion but the lower quantiles are characterised by positive coefficients and upper quantiles by negative coefficients. This is evidence of a "cushion effect" against contagion and suggests that these sectors are immune or at least that there is a decoupling in the tails regarding extreme negative movements during the credit crunch crisis (positive coefficients associated with large negative returns at lower quantiles).

CREDIT CRUNCH LEHMAN COEFFICIENTS: WEEKLY (2000–12)									
Quantile	**Staples**	**Financials**	**Industrials**	**Infotec**	**Materials**	**Telco**	**Utilities**	**Discretionary+**	**Energy+**
1%	0.0288	-0.1021*	-0.0524	-0.0144	-0.0572**	0.0429*	0.0112	0.0013	-0.0151
5%	0.0029	-0.0237	-0.0092	-0.0617	-0.0337**	-0.0050	0.0095	-0.0143	-0.0017
50%	0.0100	0.0141*	0.0068	0.0081	-0.0003	-0.0059	0.0020	0.0297**	0.0024
95%	-0.0051	0.0391**	0.0210*	0.0299	0.0114	0.0431*	0.0307	-0.0136	0.0086
99%	-0.0345	0.0240	0.0081	-0.0119	0.0490*	0.0516	0.0034	0.0146	0.0025

CREDIT CRUNCH LEHMAN COEFFICIENTS: MONTHLY (2000–12)									
Quantile	**Staples**	**Financials**	**Industrials**	**Infotec**	**Materials**	**Telco**	**Utilities**	**Discretionary+**	**Energy+**
3%	-0.1240	-0.0228	0.0244	-0.0248	0.1829***	0.0863	0.0273	0.1360	0.0874
5%	-0.1573	-0.0445	-0.0432	-0.0681	0.1519***	0.0418	-0.0103	-0.0132	0.0986
50%	0.0270	0.0533	-0.0005	0.0158	0.0184	-0.0251	-0.0190	0.0736	0.0392
95%	0.0804	0.0423	0.0866	0.1537	-0.0586	0.0465	0.1028	0.0222	-0.0017
97%	-0.0570	0.0316	0.0811	0.1188	-0.2119*	0.0765	0.0917	0.0763	-0.0025

+ In the case of the energy and discretionary SOEs sectors the weekly and monthly quantiles should be interpreted as: 3%, 5%, 50%, 95%, and 97% and 5%, 8%, 50%, 75%, and 92% respectively, since there are not enough data points to compute higher quantiles in these two sectors. ***1%; **5%; and *10% significance.

The picture is similar for the European sovereign debt (ESD) crisis (Table 3.10). In the materials, telecommunication services, utilities, and energy (weekly) and in the information technology and utilities (monthly) SOEs sector we find that the lower quantiles are characterised by positive coefficients and upper quantiles by negative coefficients, again suggesting some cushioning of extremes in the SOEs.

Table 3.10: Quantile regression estimates for for weekly and monthly holding periods – sovereign debt crisis

By generalising equation (3.1) in Section 3.3 we obtain:
$F^{-1}(\tau | SECSOE_t - COMPSEC_t) = \alpha_t(\tau) + \beta_1(\tau) MKTPREM_t + \beta_2(\tau) SMB_t + \beta_3(\tau) HML_{i,t} + \beta_4(\tau) MOM_t + \gamma_1(\tau) D_{Sub} + \gamma_2(\tau) D_{Lehman} + \gamma_3(\tau) D_{ESD} + F_{\varepsilon}^{-1}(\tau | SECSOE_t - COMPSEC_t)$ where (τ) = the 1%, 5%, 50%, 95%, and 99% for the weekly and 3%, 5%, 50%, 95%, and 97% for the monthly observations. In the materials, telecommunication services, utilities, and energy (weekly) and in the information technology and utilities (monthly) SOEs sector we reject the null of no contagion since the lower quantiles are characterised by positive coefficients and upper quantiles by negative coefficients. This is evidence of a "cushion effect" against contagion and suggests that these sectors are immune or at least that there is a decoupling in the tails regarding extreme negative movements during the sovereign debt crisis (positive coefficients associated with large negative returns at lower quantiles).

ESD COEFFICIENTS: WEEKLY (2000–12)									
Quantile	**Staples**	**Financials**	**Industrials**	**Infotec**	**Materials**	**Telco**	**Utilities**	**Discretionary+**	**Energy+**
1%	0.0278	0.0038	-0.0159	0.1290	0.0122	0.0501***	0.0637***	-0.0171	0.0317**
5%	0.0021	-0.0019	-0.0053	0.0455**	0.0080	0.0347***	0.0508***	-0.0264	0.0248**
50%	-0.0069*	-0.0086**	-0.0047	0.0003	-0.0008	-0.0069**	-0.0074*	-0.0031	-0.0016
95%	-0.0199**	-0.0150***	-0.0076	-0.0120	-0.0028	-0.0221**	-0.0404***	-0.0159	-0.0217**
99%	-0.0384	-0.0124	0.0015	-0.0712**	-0.0029	-0.0413**	-0.0811***	-0.0292	-0.0190**

ESD COEFFICIENTS: MONTHLY (2000—12)									
Quantile	**Staples**	**Financials**	**Industrials**	**Infotec**	**Materials**	**Telco**	**Utilities**	**Discretionary+**	**Energy+**
3%	-0.0119	-0.0546	-0.0331	0.1662*	0.0551	0.0141	0.1157***	0.0826	0.0254
5%	-0.0207	-0.0259	-0.0047	0.0818	0.0582	-0.0168	0.0809*	-0.0549	0.0368
50%	-0.0437**	-0.0471**	-0.0253*	0.0259	-0.0331*	-0.0386***	-0.0575**	-0.0333	-0.0015
95%	-0.0868	-0.0295	-0.0663***	0.0273	-0.2053**	-0.0528	-0.0949***	-0.1157**	-0.0339
97%	-0.1599	-0.0057	-0.0661***	0.0788	-0.2780***	-0.0397	-0.0735**	-0.1257**	-0.0353

+ In the case of the energy and discretionary SOEs sectors the weekly and monthly quantiles should be interpreted as: 3%, 5%, 50%, 95%, and 97% and 5%, 8%, 50%, 75%, and 92% respectively, since there are not enough data points to compute higher quantiles in these two sectors. ***1%; **5%; and *10% significance.

In summary, during the subprime crisis in the consumer staples, financials, industrials, consumer discretionary, and energy SOEs sectors we find evidence that (weekly) large negative excess returns *increase* but for the telecommunication services (weekly) and utilities (monthly) the lower-quantile coefficients are positive, showing a *decrease*. This decrease is evidence of a "cushion effect" during the subprime crisis. In the credit crunch crisis, weekly large negative returns *increase* in the financials and materials SOEs sectors but for the telecommunication services (short term) and materials (medium term) they *decrease*, again providing evidence of cushioning. Finally, in the ESD almost all SOE sectors are significantly affected. For materials, telecommunication services, utilities, and energy (weekly) and information technology and utilities (monthly), both very high and very low excess returns are dampened during the ESD. Telecommunications, utilities, and materials SOEs all provided some buffer against large shocks during the crisis phases.

3.4.3 Results of the cross-sectional regression

The results of the cross-sectional regressions of excess returns are summarised in Table 3.11 and Table 3.12. The incorporation of country and industry fixed effects explains most of the cross-sectional variation, and alpha (α) coefficients become insignificant. The only relevant fixed effects at the country level are for those SOEs located in Brazil and China. In the case of China the coefficient has the expected positive sign but for Brazil it is negative.

Once we allow for industry classification fixed effects we have a clearer picture of the relative performance of US benchmarks against SOEs. In both the short and the medium term the SOE sectors that outperformed their US benchmarks were financials, consumer discretionary, and consumer staples. In the remaining sectors there was no statistical evidence that SOEs outperformed. Finally, the telecommunication services sector effect was insignificant in the short term but outperformed its US benchmarks in the medium term.

Table 3.11: Cross-sectional excess returns fixed effects weekly

The results are based on the estimated coefficients from time-series regressions in a cross-sectional framework: $AVGER_{it} = c_0 + c_1 AVGER_{it-1} + c_2 AVGCAP_{it-1} + c_3 AVGBK_{it-1} + \mathbf{c_4 D_{it}} + \varepsilon_{it},\ i = 1,....n$ where $AVGER_{it}$ = the average weekly excess return net of the risk-free interest rate for each SOE and comparable US stocks that composes our sample for the period between 2000 and 2012. $AVGCAP_{it-1}$ = weekly average change in market capitalisation lagged by one period, $AVGBK_{it-1}$ = weekly average change in market-to-book value lagged by one period, and $AVGER_{it-1}$ = $AVGER_{it}$ lagged by one period. D_{it}= is the dummy variable vector that accounts for country- and industry-level fixed effects. Standard errors are adjusted using the Newey-West correction for serial correlation. ***1%; **5%; and *10% significance.

WEEKLY (2000–12)	***SOE-CROSS-SECTIONAL***	
	AVGER	**AVGER**
AVGER(-1)	0.0438	0.0312
p-value	(0.4199)	(0.5592)
AVGCAP(-1)	0.2373	0.2450
p-value	**(0.0000)** ***	**(0.0000)** ***
AVGBK(-1)	-0.0094	-0.0026
p-value	(0.8352)	(0.9536)
SOE	0.00013	
p-value	(0.5698)	
CHINA		0.00055
p-value		**(0.0425)** **
BRAZIL		-0.00089
p-value		**(0.0595)** *
INDIA		-0.00004
p-value		(0.9567)
RUSIA		-0.00054
p-value		(0.1810)

WEEKLY (2000-2012)	***SOE-CROSS SECTIONAL INDUSTRY EFFECT S***			
	FINANCIALS	**FINANCIALSSOE**	**DISCRETIONARY**	**DISCRETIONARYSOE**
AVGER	-0.0001	0.0008	0.0008	0.0024
p-value	(0.4967)	**(0.0286)** **	**(0.0000)** ***	**(0.0000)** ***

WEEKLY (2000-2012)				
	ENERGY	**ENERGYSOE**	**INFOTEC**	**INFOTECSOE**
AVGER	0.0015	0.0010	-0.0001	-0.0023
p-value	**(0.0000)** ***	(0.1459)	(0.5609)	**(0.0000)** ***

WEEKLY (2000-2012)				
	MATERIALS	**MATERIALSSOE**	**STAPLES**	**STAPLESSOE**
AVGER	0.0009	0.0001	0.0007	0.0008
p-value	**(0.0001)** ***	(0.8343)	**(0.0019)** ***	**(0.0523)** *

WEEKLY (2000-2012)				
	TELCO	**TELCOSOE**	**UTILITIES**	**UTILITIESSOE**
AVGER	-0.0011	-0.0002	0.0004	-0.0002
p-value	**(0.0036)** ***	(0.5358)	**(0.0191)** **	(0.5314)

Table 3.12: Cross-sectional excess returns fixed effects monthly

The results are based on the estimated coefficients from time-series regressions in a cross-sectional framework:
$AVGER_{it} = c_0 + c_1 AVGER_{it-1} + c_2 AVGCAP_{it-1} + c_3 AVGBK_{it-1} + \mathbf{c_4 D_{it}} + \varepsilon_{it},\ i = 1,....n$ where $AVGER_{it}$ = the average monthly excess return net of the risk-free interest rate for each SOE and comparable US stocks that composes our sample for the period between 2000 and 2012. $AVGCAP_{it-1}$ = monthly average change in market capitalisation lagged by one period, $AVGBK_{it-1}$ = monthly average change in market-to-book value lagged by one period, and $AVGER_{it-1}$ = $AVGER_{it}$ lagged by one period. $\mathbf{D_{it}}$= is the dummy variable vector that accounts for country- and industry-level fixed effects. Standard errors are adjusted using the Newey-West correction for serial correlation. ***1%; **5%; and *10% significance.

MONTHLY (2000–12)	***SOE-CROSS-SECTIONAL***	
	AVGER	**AVGER**
AVGER(-1)	0.0765	0.1379
p-value	(0.2518)	**(0.0005)** ***
AVGBK(-1)	-0.0283	0.0024
p-value	(0.5315)	**(0.0401)** **
AVGCAP(-1)	0.1740	0.0006
p-value	**(0.0024)** ***	(0.5189)
SOE	0.00116	
p-value	(0.2381)	
CHINA		0.00263
p-value		**(0.0164)** **
BRAZIL		-0.00410
p-value		**(0.0589)** *
INDIA		-0.00003
p-value		(0.9918)
RUSIA		-0.00281
p-value		(0.2049)

MONTHLY (2000-2012)	***SOE-CROSS SECTIONAL – INDUSTRY EFFECTS***			
	FINANCIALS	**FINANCIALSSOE**	**DISCRETIONARY**	**DISCRETIONARYSOE**
AVGER	-0.0004	0.0040	0.0038	0.0106
p-value	(0.6292)	**(0.0156)** **	**(0.0000)** ***	**(0.0000)** ***

MONTHLY (2000-2012)				
	ENERGY	**ENERGYSOE**	**INFOTEC**	**INFOTECSOE**
AVGER	0.0068	0.0048	-0.0008	-0.0097
p-value	**(0.0000)** ***	(0.1195)	(0.4467)	**(0.0000)** ***

MONTHLY (2000-2012)				
	MATERIALS	**MATERIALSSOE**	**STAPLES**	**STAPLESSOE**
AVGER	0.0041	0.0005	0.0036	0.0039
p-value	**(0.0000)** ***	(0.7734)	**(0.0008)** ***	**(0.0237)** **

MONTHLY (2000-2012)				
	TELCO	**TELCOSOE**	**UTILITIES**	**UTILITIESSOE**
AVGER	-0.0031	0.0015	0.0020	-0.0004
p-value	**(0.0137)** **	**(0.0937)** *	**(0.0102)** **	(0.6905)

3.5 Conclusion

By estimating a four-factor regression model of returns to BRIC economy SOEs, we can study their performance during the financial crisis and recovery period. Our model benchmarks the SOEs against comparable US firms using US market factors, including industry-sector indexes. The US market, size and value-growth factors are significant factors in models of excess returns to SOEs.

The US market premium is a significant factor for explaining excess SOEs returns for the majority of sectors with the exceptions of industrials, information technology, and energy. The US SMB factor explains part of the variation associated with extreme excess SOEs returns in the staples, financials, utilities, telecommunication services, industrials, and information technology. The HML factor also explains part of the variation of short-term extreme shocks in the financial sector. Finally, momentum is significant for large negative excess returns in the financials, materials (weekly), and telecommunication services but has a negative sign for the industrials, utilities, energy and discretionary (monthly) SOE sectors.

Additionally, by analysing dependence at extreme quantiles during crisis episodes, we observe that certain industry sectors were indeed either less or more exposed to shocks from the US than others. Our results show that there is evidence of a "cushion effect" against extreme shocks in the telecommunication, utilities, materials, information technology, and energy sectors during different episodes of the crisis. In the staples and industrial sector we observe an increased impact of the extreme returns in the subprime episode but not during the credit crunch crisis. Telecommunications, utilities, and materials SOEs all provided some buffer against large shocks during the crisis phases.

Finally, when we account for cross-sectional variation, country effects are significant and positive for China and negative for Brazil. This means that at the country level, Chinese state-owned companies were more likely to outperform their benchmarks in the US than Brazilian SOEs. For India and Russia the cross-sectional variation was statistically insignificant. At the industry level the SOEs that outperformed their US benchmarks are in the financials, consumer discretionary, and consumer staples sectors.

In the case of the financials sector, our results confirm other studies that have found that global factors can amplify shocks in the financial industry (Bagliano and Morana, 2012; Bekaert, Ehrmann, et al., 2011). Additionally, Hossain, Jain and Mitra (2013) found similar results regarding the performance of SOEs in the financial sector.

Finally, our results shed light on the role of government ownership and intervention by addressing the issue of performance of partially privatised SOEs. This "partnership" between the state and private investors can be mutually beneficial during crisis periods. Especially, due to the "cushion effect" or a decoupling in the tails, reducing extreme negative movements in certain economic sectors.

Future research in this area could take into account the different levels of government ownership and indirect intervention (subsidised loans, effect of government contracts, etc.) as well as the effects that "implicit" government guarantees have on firm performance in both developed and emerging markets.

Chapter 4 Testing for differences in sovereign spreads during the GFC using propensity-matching estimators

4.1 Introduction

A critical problem in contagion modelling is subjectivity in dating crises. The power of any test for breaks or new channels in market linkages depends on how samples are set, and different dating can lead to different results (Fry et al., 2011; Kose, 2011). Equally important is selecting a meaningful non-crisis benchmark, especially when the crisis dating approach is based on exogenously chosen events (Fry et al., 2011). Even where dates are fixed endogenously, variable selection may introduce bias (Baur, 2012).

Crisis dating methods fall into three categories: threshold-based methods where crisis dates are selected using extreme negative values at arbitrarily chosen quantiles;[27] endogenous dating models that use Markov switching regimes and/or changes in time-varying volatility for determining crisis dates;[28] and exogenous dating in which the pre-crisis and crisis periods are divided into fixed timeframes by critical events.[29] Here we employ the crisis dating used by Dungey et al. (2010), in which an event or important policy change marks the beginning or the end of a crisis, thus falling into the third category.

Our main objective and contribution is to test for and measure contagion in sovereign debt markets using an approach that is more robust to exogenous crisis dating than standard approaches. We use propensity matching combined with an average treatment effect on the treated (ATET) method to correct possible sample

[27] Examples in this category are: Bae, Karolyi, and Stulz (2003), Longin, and Solnik (2001), Kaminsky, Lizondo, and Reinhart (1998) and Eichengreen, Rose, and Wyplosz (1996).

[28] See for example Ang and Bekaert (2002), Dungey, Milunovich, and Thorp (2010a) and Phillips and Yu (2011).

[29] This is the most common approach, and the body of literature is too large to be included. A comprehensive survey of the literature can be found in Fry et al. (2011).

selection effects. Propensity-matching methods borrow from the methods of randomised controlled trials: at the first stage, general factor models, including crisis dummies, are fitted to the whole sample; then a set of non-crisis observations most closely matching the factor values of the crisis sample observations are drawn, building an artificial but matching "control" sample; and finally, the crisis and artificial non-crisis samples are compared in formal tests of differences in spreads. In this way, by allowing our crisis observations to act as "treated" units, we can test whether the difference in spreads versus our "non-treated" benchmark is statistically significant. We apply this method to test for contagion in sovereign debt markets during the recent crises.

Our sample includes debt securities from 43 countries grouped into nine different portfolios: all economies debt, developed economies debt, emerging economies debt, Euro-currency debt, US dollar-denominated debt, local-currency debt, local-currency developed economy debt, local-currency developing economy debt, and the troubled European countries: Portugal, Ireland, Italy, Greece, and Spain (PIIGS).

Although there is no agreement on a single standardised factor model for sovereign spreads, there is consensus that country-specific fundamentals have been a major determinant of the variability of sovereign spreads during the financial crisis (GFC). In fact, several studies concur that prior to and during the first stage of the GFC in 2007–08, global risk aversion was driving sovereign spreads (Caceres et al., 2010; Sgherri and Zoli, 2009), but from 2009 onwards, country-specific fundamentals became dominant. The perceived fragility of a country's financial sector and its potential to deplete public finances, signs of weak macroeconomic fundamentals, and changes in trade variables became important to explaining differences in the sovereign spreads of Eurozone countries from the beginning of the global recession and into the recent sovereign debt crisis (Mody, 2009; Schuknecht, Von Hagen, and Wolswijk, 2009; Arghyrou and Kontonikas, 2012).

In the case of emerging-market bonds, country fundamentals and proxies for risk aversion and liquidity are also major determinants of emerging markets spread variation (González-Rozada and Yeyati, 2008; Hilscher and Nosbusch, 2010; Remolona et al., 2007). This increase in the importance of country-specific factors during the GFC stands in sharp contrast to previous crises in emerging countries, where spreads were driven mainly by global factors (Martinez et al., 2013; Mauro et al., 2002).[30] There are many possible sources of influence on sovereign spreads, and it is natural to ask what factors – global, country-specific or latent – are the main drivers behind the changes in sovereign spreads.

Studies of contagion offer a taxonomy of transmission channels that can be used to categorise factors affecting spreads. Dungey and Martin (2007) classify the transmission channels into three categories: common or market shocks; country-specific shocks; and latent or idiosyncratic shocks. In this paper we use the definitions of Giordano et al. (2013), in which common markets shocks are referred as "shift contagion", country-specific transmissions as "wake-up" contagion, and latent factor transmissions as "pure contagion". This three-way classification gives an economic interpretation of the factor model estimated here and allows an analysis of channels of volatility transmissions across sovereign debt markets. The main contribution in this paper is that we propose a novel framework to test for differences in spreads that correct biases found in conventional exogenous dating methods using

[30] Other studies attribute the increase in global liquidity to the fall of emerging-market spreads and a shift from common factors to specific factors during the GFC (Eichengreen et al., 2012; Hartelius et al., 2008). In the case of sovereign credit default swaps (CDS), which are common proxies for sovereign spreads, unobservable factors and risk aversion account for a large part of the observed variation (Coudert and Gex, 2008; Longstaff et al., 2011).

propensity-matching estimators. We show that the differences in spreads between crisis and non-crisis periods when obtained with traditional exogenous dating methods are grossly underestimated or overestimated when compared with results obtained with matching estimators. We present evidence that the portfolio of local-currency emerging-market debt did not exhibit any significant difference in spreads during the GFC as a whole, even under robust specifications, and that the earlier phases of the GFC were not as contagious as previously thought, at least in the case of sovereign debt.

This paper is divided as follows: Section 4.2 sets out data sources and choices of variables for base regression models; section 4.3 outlines the proposed empirical model for measuring the differences between non-crisis and crisis conditions, Section 4.4 contains the summary of results, and Section 4.5 concludes.

4.2 Common determinants of sovereign spreads

4.2.1 Variables and data description

In order to observe the effects that country-specific and market factors had in sovereign spreads during the Global Financial Crisis (GFC) we build factor models consistent with existing studies. However, instead of credit default swap (CDS) spreads, we model spreads of sovereign zero coupon bonds. The main reason for using actual spreads rather than CDS spreads is that CDS are priced using a risk-neutral framework, therefore default probabilities for CDS appear much higher than those inferred from historical bond prices (Hull et al., 2012). By using actual spreads we address the issue of upward bias in the implied default probabilities described in Section 4.3.

For this paper, we used as our proxies for sovereign yields the Bloomberg fair market value zero coupon denominated sovereign bond curves (FMCZCB) for 43 countries (of which 23 are developed and 20 are emerging markets). These curves have the distinctive feature that they are derived from actual bond prices and give a good approximation of what would be the theoretical price of other maturities that are not traded. The zero coupon curves are calculated for a different range of maturities using exactly the same base model created by Bloomberg, hence it is easy to aggregate prices into a country portfolio as well as observe differences in spreads relative to the US FMCZCB for different maturities. This feature allows aggregation by market weights, better measuring the actual effect of a country's sovereign spread relative to its economic importance. We compute the weights of sovereign debt securities for any country *(i)* using the following formula:

$$w_{\mathrm{j,i}} = \frac{v_{j.i}}{\sum_{j=1}^{n} v_{j.i}} \tag{4.1}$$

where $v_{j.i}$ is the total currency value of a sovereign bond with a maturity *(j=1,2..n)* in a country (*i*), $\sum_{j=1}^{n} v_{j.i}$ is the sum of the total currency value of all *n* issues of sovereign bonds in country *(i)*, and $w_{\mathrm{j,i}}$ is the percentage (%) weight of a sovereign bond with a maturity *(j)* of the total currency value outstanding of all issues of sovereign bonds in a country *(i)*. Also, we compute the value-weighted theoretical yield for sovereign debt any given country (*i*) at time (*t*) using the following formula:

$$Y_{\mathrm{i},t} = \sum_{j=1}^{n} w_{j,i}\, \mathrm{y}_{i,\mathrm{j},t} \tag{4.2}$$

where $y_{i,j,t}$ =is the yield at time (t) for bonds of any country ($i=1,\dots,43$) of maturity (j). Notice that $w_{j,i}$ is kept constant[31] at all times (t). $Y_{i,t}$ is the proxy for the market-weighted theoretical yield for a country (i). Another distinctive advantage of zero coupon yields is that it is easy to compute the weighted average theoretical duration of currently traded issues. Therefore, the weighted average duration for a set of different maturity country yields is given by the following formula:

$$\mathrm{Dur}_i = \sum_{j=1}^{n} w_{j,i} dur_{j,\mathrm{i}} \tag{4.3}$$

where Dur_i is the weighted average duration of country (i) in years and and $dur_{j,\mathrm{i}}$ is the years to maturity ($j=1,2..n$) of a sovereign bond in a country (i). Finally the observed spread for any country (i) is given by the following formula:

$$spread_{i,t} = Y_{i,t} - Y_{us,\mathrm{t}\approx\mathrm{Dur}_i} \tag{4.4}$$

where $Y_{i,t}$ = is the proxy for the market-weighted theoretical yield and $Y_{us,\mathrm{t}\approx\mathrm{D}_c}$ is the yield of the USD FMCZCB with closest maturity to the duration obtained in equation (4.3).

Using daily data for the FMCZCB from January 3, 2000, to May 31, 2013, we compute the monthly average yield for each country and calculate the theoretical market-weighted spread using equations (4.1) to (4.4) and aggregate them at the portfolio level for nine (9) groups: an "all countries" portfolio, which includes the 43 countries in the sample; "developed" and "emerging" countries portfolios, which are

[31]In order to compute the weights, we use the last reported total issued amount outstanding as of May 31, 2013, since there is no longitudinal data source on amount issued, just the snapshot at the collection time. This is a data limitation problem which in our opinion does not affect the results much since most countries tend to rollover maturing debt with new issues of similar amounts.

divided according to MSCI classification before the European sovereign debt (ESD) crisis in which Greece is considered a developed country and not an emerging market as the post-ESD reclassification; a "Euro" portfolio, which includes all countries from the Eurozone that issue their debt in euros; a "USD" portfolio, which includes the countries that issue debt in US dollars; a "Local" portfolio, which includes all those countries that issue their debt in local currency; and finally "local developed" and "local emerging" countries portfolios for sovereigns that issue debt in their home currency; and the troubled countries' (Portugal, Ireland, Italy, Greece and Spain (PIIGS)) portfolio. Table 4.1 reports descriptive statistics for the portfolio spreads obtained from the sample. In order to aggregate spreads into the portfolios we apply equation (4.5):

$$SPREAD_{p,t} = \left(\frac{V_{i,p}}{\sum_{i=1}^{p} V_{i,p}} \right) spread_{i,\mathrm{p,t}} \qquad (4.5)$$

Where $spread_{i,\mathrm{p,t}}$ is the spread from equation (4) for a country *(i)* that is part of portfolio *(p)* at time *(t)*, $V_{i,p}$ = the total value outstanding of a member country *(i)* of portfolio *(p)* converted to US dollars in case of issues in euros or local currency using the exchange rate of May 31, 2013, for the same reasons explained in footnote 31, $\sum_{i=1}^{p} V_{i,p}$ = the sum of the total value of sovereign bonds outstanding for all the countries *(i)* that are part of a specific portfolio *(p)*, and $SPREAD_{p,t}$ is the market value weighted spread of portfolio *(p)* at time *(t)*.

Table 4.1: Descriptive statistics sovereign spreads (monthly)

This table reports the descriptive statistics of the market-weighted spreads over the US zero coupon yield of the same maturity (equation (4.5)) as reported by Bloomberg (FMCZCB) from January 1, 2000, to May 31, 2013, for the 43 countries in our sample. The results are reported in basis points on a monthly basis. A negative spread sign implies that the spread in that portfolio was lower on average than that in the US during the period of observation. The spreads are aggregated into nine (9) market-weighted portfolios. The countries whose debt comprise each portfolio are below their respective columns.

	All countries	**Developed**	**Emerging**	**EUR**	**USD**	**LOCAL**	**Local developed**	**Local emerging**	**PIIGS**
Mean	110.51	-9.37	273.42	23.66	351.78	76.62	-22.30	224.70	95.93
Median	37.75	-38.53	222.57	-36.50	247.46	36.62	-35.90	227.41	-22.24
Maximum	4147.66	4147.66	2129.80	4147.66	2129.80	1010.10	1519.81	1010.10	4147.66
Minimum	-569.53	-569.53	-369.39	-265.31	0.00	-569.53	-569.53	-369.39	-265.31
Std. dev.	287.78	236.71	270.48	285.55	308.29	232.80	195.26	226.69	417.91
Skewness	2.85	6.70	1.48	7.26	1.78	0.53	1.77	0.34	4.98
Kurtosis	25.41	91.26	7.74	80.82	7.58	3.32	13.66	2.77	38.21
Observations	6407	3691	2716	1881	1151	3375	2093	1443	793
Colombia	Japan	Austria	Colombia	Austria	Colombia	Chile	Australia	Chile	Portugal
Brazil	Malaysia	Australia	Brazil	Belgium	Brazil	Peru	Canada	Peru	Italy
Mexico	Netherlands	Belgium	Mexico	Finland	Mexico	Australia	Denmark	Bulgaria	Ireland
Venezuela	New Zealand	Canada	Venezuela	France	Venezuela	Bulgaria	Hong Kong	Czech Republic	Greece
Chile	Norway	Denmark	Chile	Germany	Indonesia	Canada	Japan	Hungary	Spain
Peru	Philippines	Finland	Peru	Greece	Philippines	Czech Republic	New Zealand	India	
Austria	Poland	France	Bulgaria	Ireland	Russia	Denmark	Norway	Malaysia	
Australia	Portugal	Germany	Czech Republic	Italy	Slovakia	Hong Kong	Singapore	Poland	
Belgium	Romania	Greece	Hungary	Netherlands	Turkey	Hungary	Korea	Romania	
Bulgaria	Russia	Hong Kong	India	Portugal		India	Sweden	South Africa	
Canada	Singapore	Ireland	Indonesia	Slovakia		Japan	Switzerland	Thailand	
Czech Republic	Slovakia	Italy	Malaysia	Singapore		Malaysia	United Kingdom		
Denmark	South Africa	Japan	Philippines	Spain		New Zealand			
Finland	Korea	Netherlands	Poland			Norway			
France	Spain	New Zealand	Romania			Poland			
Germany	Sweden	Norway	Russia			Romania			
Greece	Switzerland	Portugal	Slovakia			Singapore			
Hong Kong	Thailand	Singapore	South Africa			South Africa			
Hungary	Turkey	Korea	Thailand			Korea			
India	United Kingdom	Spain	Turkey			Sweden			
Indonesia		Sweden				Switzerland			
Ireland		Switzerland				Thailand			
Italy		United Kingdom				United Kingdom			

From the descriptive statistics we can observe that many of the portfolios have high kurtosis and skewness, which reflects a large degree of heterogeneity in the sample. As expected, the PIIGS portfolio exhibits higher volatility in wake of the European sovereign debt crisis, and the local-currency developed countries portfolio exhibits the lowest volatility. The reason that some portfolios exhibit negative spreads can be explained by the fact that some developed countries like the UK, Germany, and Japan had lower nominal rates than the US on average during the period under study.

Our choice of explanatory variables for common factor or “shift contagion” includes the global and US equity premiums, the European bond index, a regional bond index, as well as the global risk-aversion index. The global equity premium and US equity premiums are proxied by the S&P Global and S&P 500 indexes net of exchange rate variation and the risk-free rate. In order to account for the effect of the regional bond prices, we use the approach employed by Longstaff et al. (2011) by including the returns to the regional bond portfolio excluding the country under observation. For the European bond index we use the EFFA, which is a Bloomberg market-weighted index that includes all the Eurozone government debt with a maturity longer than one year. Finally, the Chicago volatility index (VIX) proxies for global risk aversion. All the data for common or “shift-contagion” proxies were extracted from Bloomberg. In the case of the explanatory variables for country-specific determinants or “wake-up contagion”, the local premium is represented by the changes of the local stock market of each country, adjusted by the domestic currency/USD exchange rate. We also deduct the risk-free rate, here proxied by the US Treasury zero coupon yield of similar maturity to the constructed bond portfolio for each country. We also include a set of macroeconomic variables commonly used in other studies and that are explained in detail in Section 4.2.2. All data for the country-specific determinants were extracted from the IMF statistics module in Bloomberg in order to guarantee harmonisation among the variables, except the growth rate of GDP per capita provided by the World Bank.

4.2.2 Factor model

To measure the effect of country-specific determinants in sovereign spreads we begin with a panel data model. Martinez et al. (2013) argue that panel data models can deal with the cross-sectional heterogeneity and time effects that are present in macroeconomic data. Baur and Fry (2009) argue that in a panel data model with common factors, significant time fixed effects capture the latent or pure contagion factor. Most of the studies[32] that attempt to explain the behaviour of spreads use macroeconomic variables as country-specific factors; here we follow the specification proposed by Giordano, Pericoli, and Tommasino (2013), where:

$$spread_{i,t} = \alpha_o + \alpha_1 spread_{i,t-1} + \beta_{i,t} Z_{i,t} + \beta_{i,t} F_t + \gamma_o D_{c,t} + \gamma_1 Z_{i,t} D_{c,t} + \gamma_2 F_t D_{c,t} + \varepsilon_{i,t} \quad (4.6)$$

Where α_0 is the common intercept; $Z_{i,t}$ is a vector of country-specific factors which in our model are the exchange rate, total debt to GDP ratio, investment to GDP ratio, external debt to exports ratio, GDP per capita growth, reserves, and the local equity premium;[33] and F_t is a vector of common factors which in our model are the global equity premium, US equity premium, the regional bond portfolio and the global risk

[32] Recent examples of studies that apply panel techniques for explaining sovereign spreads using country-specific determinants can be found in Balazs and Ivaschenko (2013), Beirne and Fratzscher (2013), Aizenman et al. (2012), and Hilscher and Nosbusch (2010).

[33] The amount of literature about country-specific proxies is beyond the scope of this paper, but some classic examples can be found in Eichengreen and Mody (1998), Boehmer and Megginson (1990), Edwards (1984), Edwards and Levy Yeyati (2005), Dittmar and Yuan (2008), and Berg et al. (2005) just to mention a few.

aversion,[34] and the subscripts i and t stand for country and month respectively. The expected sign of the coefficients of both country-specific ($Z_{i,t}$) and common (F_t) factors are summarised in Table 4.2:

[34] Some examples of using the common factors and proxies of global risk aversion stock can be found in Coudert and Gex (2008), Longstaff et al. (2011), and Dahiya (1997).

Table 4.2: Determinants and their relation with sovereign spreads

Country-specific factors	***Common factors***
1) The exchange rate is expected to have a positive (+) coefficient since depreciations are associated with weaker economic conditions and higher spreads.	1) The global market equity premium acts as a transmission mechanism of global conditions and is expected to have a negative (-) coefficient if there is a strong global outlook; under weak global prospects it is expected to have a positive (+) sign.
2) The ratio of debt to GDP (Debt/GDP) is expected to have a positive (+) coefficient. An increase in this ratio implies an increase in the probability of default. Consequently, the creditors would require a higher spread in order to compensate for this additional risk.	2) The European bond index acts as a plausible transmission mechanism of global conditions during the European sovereign debt crisis and can either have a negative (-) or positive (+) coefficient depending on the region of analysis.
3) The ratio of investment to GDP (Investment/GDP) whether the sign is positive (+) or negative (-) is still an ongoing debate. A higher investment ratio can be tied to future GDP growth and better economic perspectives, so if this is the case the sign of the coefficient is expected to be negative. However, a higher investment ratio can also be financed by increasing public debt and if this is the case, the coefficient is expected to have a positive sign.	3) The US equity premium acts as a plausible transmission mechanism of global conditions during the subprime and Lehman Brothers crisis episodes and can either have a negative (-) or positive (+) coefficient depending on the region of analysis.
4) The ratio of debt to exports (Debt/Exports) acts as a proxy for debt service and liquidity. A higher ratio is related to lower liquidity and a greater strain on available resources to meet future debt servicing obligations, so for this variable we expect a positive (+) sign.	4) The regional bond portfolio acts as a plausible transmission mechanism of regional conditions during all crisis episodes and can either have a negative (-) or positive (+) coefficient depending on the region of analysis.
5) GDP per capita is expected to be negatively (-) correlated with spreads. A positive increase in GDP per capita can be interpreted as a proxy for country development and enhanced terms of credit due to future expectations of GDP growth.	5) The global risk-aversion index acts as the proxy for aggregate risk aversion during all crisis episodes and is expected to have a (+) sign during crisis periods as more risk-adverse investors demand a higher spread.
6) The ratio of current account to GDP (Current Account/GDP) acts as a proxy for liquidity. A negative ratio represents a deficit and less liquidity to meet future obligations, so in this case we expect a negative (-) sign. A positive ratio represents a surplus and more liquidity to meet future obligations, so in this case we expect a positive (+) sign.	
7) Reserves to GDP ratio (Reserves) are inversely correlated with spreads so for higher reserves we expect a negative (-) sign. The higher the foreign currency reserve the more likely is the country to meets its obligations. Imports behave in exactly the opposite way.	
8) The local equity premium is expected to have a negative (-) coefficient. An increase in the local stock market return is related to the perception of strong economic growth.	

In order to test for a specific channel of transmission during a crisis period we look at the significance of the γ coefficients from equation (4.6) during the crisis periods ($D_{c,t}$) with the country-specific and common factors. Additionally, the latent factor is represented by ($\gamma_o D_{c,t}$), in this case a significant γ_o can be interpreted as "pure contagion" or a latent factor that is neither related to the change or level in country fundamentals or common factors, but possibly attributable to unobservable factors (G. Calvo, 1988; G. A. Calvo and Mendoza, 2000). A significant γ_1 can be interpreted as "wake-up contagion" or a change in country-specific factors that leads investors to reassess their investment position in one country based on similarities of country factor fundamentals in crisis countries (Goldstein, 1998). Finally, a significant γ_2 can be interpreted as "shift contagion" or increases in the correlation of a global factor with a set of countries or regions during a crisis period (Bekaert, Ehrmann, et al., 2011; Forbes and Rigobon, 2002).

Crisis indicator variables take the value of one in each of the three different phases of the GFC between July 26, 2007, and May 17, 2012. We name and date the crisis phases as follows: The first-phase "subprime" crisis (D_{Sub}) begins July 26, 2007, which was the day the Dow Jones recorded a significant large loss in response to bad news from mortgage lender Countrywide Financial. At this point, the market processed news of "difficult conditions" in the subprime market following Countrywide Financial Corporation's SEC filing on July 24. The beginning of the "credit crunch" crisis (D_{Credit}) is generally dated from the time Lehman Brothers filed for bankruptcy on September 15, 2008. The European sovereign debt crisis (D_{ESD}) we date from October 22, 2009, when Fitch first downgraded and reported a negative outlook for Greek sovereign debt, until May 17, 2012, when the same agency upgraded it again from a default rating due to the compromise reached by the Greek government with the European monetary authorities.[35]

Finally, in order to deal with possible misspecification issues present in the base regression we use a two-step process to select the most relevant variables. In the first step we estimate a stepwise (backward and forward) panel OLS regression to eliminate those variables that are statistically insignificant for each of the nine (9) portfolios groups mentioned in Subsection 4.2.1. In the second step we estimate a second OLS regression with the obtained variables, but this time including country fixed effects or country heterogeneity and robust (heteroskedasticity consistent) standard errors. The stepwise procedure in variable selection has been used by Carrieri, Errunza, and Hogan (2007) for selecting variables for market integration, and the robust error specification has been used by Longstaff et al. (2011) for analysing CDS sovereign spreads. This procedure is repeated for each of the nine (9) portfolios for a total of nine regressions. In Table 4.3 we report the results obtained for the whole period of the GFC.

[35] The key dates for the subprime and credit crunch crises were taken from the financial turmoil timeline chart from the Federal Reserve Bank of St Louis (http://timeline.stlouisfed.org/pdf/CrisisTimeline.pdf) and for the European sovereign debt crisis from the credit rating function in Bloomberg. There are other studies that use similar dates for the credit crisis and place the subprime around the same period, including Frank and Hesse (2009), Dooley and Hutchison (2009), and Felices and Wieladek (2012).

Table 4.3: Panel regression estimates

This table reports the results of the panel regression: $spread_{i,t} = \alpha_o + \alpha_1 spread_{i,t-1} + \beta_{i,t} Z_{i,t} + \beta_{i,t} F_t + \gamma_o D_{\mathrm{GFC},t} + \gamma_1 Z_{i,t} D_{\mathrm{GFC},t} + \gamma_2 F_t D_{\mathrm{GFC},t} + \varepsilon_{i,t}$ where $spread_{i,t}$ is the spread of any country i at time t from January 2000 to May 2013, α_0 is the common intercept, $Z_{i,t}$ is a set of country-specific factors which in our model are the exchange rate, total debt to GDP ratio, investment to GDP ratio, external debt to exports ratio, GDP per capita growth, current account to GDP, imports, reserves, and local equity premium. F_t is a set of common factors which in our model are the global equity premium, US equity premium, the regional bond portfolio and the global risk aversion, and the subscripts $_i$ and $_t$ stand for country and month respectively. D_{Sub} represents the dummy for the months of the "subprime" phase of the GFC from July 2007 until August 2008. D_{Credit} represents the dummy for the months of the "credit crunch" crisis from September 2008 until September 2009. D_{ESD} represents the dummy for the months of the European sovereign debt crisis from October 2009 until May 2012. γ_0 represents "pure contagion", γ_1 represents "wake-up contagion" and γ_2 represents the "common-factor contagion". ***1%; **5%; and *10% significance. Macroeconomic variables as GDP are issued quarterly so we keep them constant on a monthly bais as is common in the literature detailed in footnote 33. In order to deal with endogeneity and possible misspecification issues present in the base regression we run a two-step process to select the most relevant variables. In the first step we run a stepwise panel OLS regression with backward and forward inclusion to drop those variables that are statistically insignificant for each of the nine (9) portfolios groups mentioned in Subsection 4.2.1. In the second step we run a second OLS regression with the remaining variables, but adjusting for cross-sectional effects and robust (heteroskedasticity consistent) standard errors.

	All countries	**Developed**	**Emerging**	**EUR**	**USD**	**LOCAL**	**Local developed**	**Local emerging**	**PIIGS**
Spread (t-1)	0.9618***	0.9572***	0.9617***	0.9455***	0.9586***	0.9534***	0.9662***	0.9472***	0.9397***
	(0.0293)	(0.0556)	(0.0127)	(0.0561)	(0.0188)	(0.0071)	(0.0161)	(0.0100)	(0.0523)
Country-specific factors									
Exchange rate	0.0134***		0.0195***		0.0206**	0.0033		0.0094	
	(0.0042)		(0.0058)		(0.0084)	(0.0024)		(0.0061)	
Debt/GDP	-0.0007	-0.0011					-0.0007		
	(0.0012)	(0.0029)					(0.0006)		
Investment/GDP									
Debt/exports	0.0010		0.0013**		0.0018**		-0.0006		
	(0.0008)		(0.0005)		(0.0007)		(0.0005)		
GDP per capita growth	0.0007	0.0027		0.0115		0.0032	0.0046**	0.0040	
	(0.0028)	(0.0040)		(0.0091)		(0.0021)	(0.0021)	(0.0029)	
Current account/GDP									

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
Reserves						-0.0001		-0.0020	
						(0.0003)		(0.0014)	
Local equity premium	-0.0155***	-0.0028*	-0.0132***		-0.0162***	-0.0158***	-0.0153**	-0.0091***	
	(0.0036)	(0.0016)	(0.0025)		(0.0041)	(0.0050)	(0.0061)	(0.0023)	
Common factors									
Global equity premium	-0.0069*					-0.0098*	-0.0131**		
	(0.0035)					(0.0052)	(0.0062)		
European bond index	0.0436***	-0.0951***	0.0430***		0.0265*	0.0549***		0.0582***	
	(0.0115)	(0.0299)	(0.0125)		(0.0147)	(0.0103)		(0.0115)	
US equity premium			-0.0142**		-0.0249**	0.0065	0.0127**		
			(0.0070)		(0.0112)	(0.0056)	(0.0064)		
Regional bond portfolio	-0.1405	1.1907***		0.3234***		-0.1089	0.3935***		0.4395***
	(0.1090)	(0.2935)		(0.1106)		(0.0917)	(0.1040)		(0.1490)
Global risk aversion	-0.0021**		-0.0023*		-0.0032*	-0.0012			
	(0.0009)		(0.0012)		(0.0017)	(0.0007)			

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
Pure contagion									
Dummy subprime (D_{Sub})	0.0622	0.0704	0.0230	0.1131**	-0.1721*	0.1020***	0.0478*	0.1447***	0.1384***
	(0.0432)	(0.0554)	(0.0792)	(0.0476)	(0.1024)	(0.0321)	(0.0251)	(0.0441)	(0.0432)
Dummy credit (D_{Credit})	-0.0875	-0.1126*	0.0030	-0.0063	-3.0413***	-0.0972	-0.1026***	0.1363	-0.0090
	(0.0693)	(0.0589)	(0.0986)	(0.0719)	(0.4359)	(0.0591)	(0.0316)	(0.3443)	(0.0857)
Dummy ESD (D_{ESD})	0.0083	0.3985***	0.0628*	1.2750***	-0.0602	0.0945***	0.0250	0.1763*	5.4985
	(0.0674)	(0.1542)	(0.0356)	(0.3609)	(0.0533)	(0.0262)	(0.0371)	(0.0912)	(3.8775)

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
Wake-up contagion									
Exchange rate x D_{Sub}	-0.0129**		-0.0240**			-0.0120*		-0.0252*	
	(0.0052)		(0.0108)			(0.0068)		(0.0131)	
Exchange rate x D_{Credit}	0.0069			0.0213					
	(0.0100)			(0.0168)					
Exchange rate x D_{ESD}	-0.0187***		-0.0288***		-0.0351***			-0.0165*	-0.0767
	(0.0059)		(0.0069)		(0.0092)			(0.0094)	(0.0623)
Debt/GDP x D_{Sub}									
Debt/GDP x D_{Credit}			-0.0040			0.0032			
			(0.0033)			(0.0021)			
Debt/GDP x D_{ESD}	0.0018*						0.0044***		
	(0.0010)						(0.0013)		
Investment/GDP x D_{Sub}									
Investment/GDP x D_{Credit}					0.1071***			-0.0213	
					(0.0120)			(0.0179)	
Investment/GDP x D_{ESD}		-0.0143**		-0.0526***				-0.0060*	-0.1694*
		(0.0069)		(0.0171)				(0.0035)	(0.0884)
Debt/exports x D_{Sub}									
Debt/exports x D_{Credit}			0.0025		0.0056**	-0.0029*		0.0046	
			(0.0031)		(0.0028)	(0.0017)		(0.0040)	
Debt/exports x D_{ESD}		0.0013					-0.0043***	0.0030**	-0.0081
		(0.0012)					(0.0015)	(0.0013)	(0.0079)
GDP per capita growth x D_{Sub}	0.0035		0.0074		0.0159**				
	(0.0039)		(0.0059)		(0.0077)				
GDP per capita growth x D_{Credit}	0.0075				-0.0344**				
	(0.0114)				(0.0156)				

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
GDP per capita growth x D_{ESD}									0.0649
									(0.0543)
Current account/GDP x D_{Sub}									
Current account/GDP x D_{Credit}			-0.0083**			-0.0062		-0.0095*	
			(0.0036)			(0.0039)		(0.0050)	
Current account/GDP x D_{ESD}	-0.0092**	-0.0208***		-0.0529***			-0.0154***		-0.1720**
	(0.0037)	(0.0055)		(0.0116)			(0.0045)		(0.0854)
Reserves x D_{Sub}									
Reserves x D_{Credit}			0.0071**		0.0530***				
			(0.0031)		(0.0189)				
Reserves x D_{ESD}		0.0041***						0.0027***	
		(0.0016)						(0.0010)	
Imports/GDP x D_{Sub}									
Imports/GDP x D_{Credit}			-0.0399					0.0956***	
			(0.0323)					(0.0357)	
Imports/GDP x D_{ESD}		-0.0066*							-0.9423
		(0.0035)							(0.9688)
Local equity premium x D_{Sub}							0.0354**	-0.0217***	
							(0.0146)	(0.0075)	
Local equity premium x D_{Credit}	0.0168***		0.0100*		0.0302***	0.0244***	0.0357***		
	(0.0059)		(0.0057)		(0.0098)	(0.0050)	(0.0059)		
Local equity premium x D_{ESD}	-0.0073		-0.0170**		-0.0652**		0.0280	-0.0354***	0.1862
	(0.0075)		(0.0083)		(0.0269)		(0.0199)	(0.0114)	(0.1470)

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
Common-factor contagion									
Global equity premium x D_{Sub}						0.0044	0.0403**	-0.0203***	
						(0.0032)	(0.0164)	(0.0078)	
Global equity premium x D_{Credit}	0.0174***					0.0312***	0.0354***		
	(0.0059)					(0.0046)	(0.0061)		
Global equity premium x D_{ESD}		0.0176	-0.0173*	0.0242	-0.0690**		0.0345*	-0.0340***	0.2115
		(0.0146)	(0.0093)	(0.0233)	(0.0288)		(0.0197)	(0.0120)	(0.1579)
European bond index x D_{Sub}						-0.0331	-0.1892***	-0.0620**	
						(0.0205)	(0.0167)	(0.0280)	
European bond index x D_{Credit}							-0.0855***		
							(0.0302)		
European bond index x D_{ESD}	-0.0536***	-0.0454	-0.0432**	-0.1820**		-0.0405***	-0.0334**	-0.0603***	-0.5840**
	(0.0202)	(0.0288)	(0.0177)	(0.0880)		(0.0131)	(0.0133)	(0.0169)	(0.2481)
US equity premium x D_{Sub}							-0.0398**		
							(0.0167)		
US equity premium x D_{Credit}		0.0302***			-0.0460			0.0328***	
		(0.0079)			(0.0322)			(0.0079)	
US equity premium x D_{ESD}	-0.0123	-0.0311*		-0.0553*	0.0630*	-0.0131*	-0.0408*		-0.1819
	(0.0091)	(0.0176)		(0.0308)	(0.0325)	(0.0069)	(0.0210)		(0.1508)
Regional bond index x D_{Sub}							1.7600***		
							(0.4517)		
Regional bond index x D_{Credit}	0.4304***	0.8249***	0.6604***		0.5272*	0.3435***	1.5923***	0.2269	

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
	(0.1604)	(0.1900)	(0.2188)		(0.3143)	(0.1075)	(0.3140)	(0.1501)	
Regional bond index x D_{ESD}				0.9210					5.0835
				(0.9070)					(3.2470)
Global risk aversion x D_{Sub}	0.0020					0.0034*			
	(0.0017)					(0.0018)			
Global risk aversion x D_{Credit}	0.0062***		0.0117**		0.0130	0.0052**	0.0040**	0.0047***	
	(0.0023)		(0.0056)		(0.0097)	(0.0020)	(0.0017)	(0.0017)	
Global risk aversion x D_{ESD}		-0.0033		-0.0073				-0.0037*	
		(0.0035)		(0.0068)				(0.0021)	
Adjusted R^2	0.9742	0.9637	0.9719	0.9552	0.9671	0.9886	0.9874	0.9803	0.9537
Number of observations	6325	3646	2679	1869	1126	3330	2058	1432	788

In the case of the Eurozone and the PIIGS portfolios, the regional bond portfolio is the most relevant common factor in explaining spread variation. In addition, when we allow for changes in channels of transmission using crisis dummies, we find evidence of latent factor contagion in the subprime and ESD crises. In the case of wake-up contagion in the Eurozone and PIIGS, the most significant determinants are the investment to GDP ratio and the current account deficit to GDP, which are liquidity related. Common factor contagion in both portfolios is explained by the European bond index and US equity premium (Eurozone), and just the European bond index in the case of the PIIGS. Beirne and Fratzscher (2013) reached a similar conclusion using data at the individual country level, but in our case we found no evidence that fundamental or "wake-up" contagion had a greater impact than common factor contagion at least at the aggregate level.

When we compare the common characteristics of the emerging- and developed-market portfolios, we observe that country-specific factors such as the debt to GDP ratio and the exchange rate are more significant in emerging markets. In the case of common factors, the impact in spread variation is larger in developed countries. When we test for "wake-up" contagion in the different phases of the crisis, both the developed- and emerging-market portfolio are sensitive to changes in fundamentals related to liquidity during the ESD phase (developed) and the credit phase (emerging). Giordano et al. (2013) argued that this increase in fundamentals significance in developed countries during the ESD was due to bad news originating from the PIIGS that led investors to closely monitor country-specific liquidity proxies in other countries.

Finally, when we control for the currency denomination of debt (USD, local, local developed, and local emerging) an interesting result is that the exchange rate is a significant common factor in the USD-denominated portfolio but not in the other three. In the case of the local developed and local emerging currency portfolios the determinants behave similarly to the emerging and developed counterparts without adjusting for currency. However, an important difference is that in the case of the local developed currency portfolio, common factor contagion by multiple channels during all phases is the key source of spread variation. In the case of the local emerging currency portfolio there is "wake-up" contagion during the credit phase and common factor contagion during the ESD crisis. The global market premium as a common factor source is just relevant during the subprime phase. Similar to Balazs and Ivaschenko (2013), we found evidence that global risk aversion becomes a significant factor during the crisis, but in the case of the local developed and local emerging currency portfolio this significance seems to be more related to common factors than to country-specific fundamentals or "wake-up" contagion.

In the case of additional channels of transmission there is a "pure contagion" latent factor in most of the portfolios, but we hypothesise that the latent factors can be explained by a global bond portfolio that is not accounted for in the model, since we observe that once all countries are aggregated into one portfolio the latent factor contagion disappears. In the case of the USD and local developed portfolios the latent factor, albeit significant, exhibits a negative sign which can be evidence of "positive contagion" as defined by Baur and Fry (2009). Finally, the local equity premium is the most common significant characteristic across all portfolios with the exception of the Eurozone and PIIGS portfolios.

4.3 Proposed empirical method for testing differences in spreads during the GFC

At the next stage we use factors identified as significant from the estimation of equation (4.6) for each of the nine portfolios, and we can obtain the implied probability of being in a crisis period using the following logit form:

$$\Pr(D_{\mathrm{GFC},t} = 1 | X_{i,t}) = (1 + \exp(-\alpha_o - \beta_{i,t} X_{i,t}))^{-1} \quad (4.7)$$

where $D_{\mathrm{GFC},t} =$ is an indicator function denoting the Global Financial Crisis (GFC), which encompasses all the three phases; $X_{i,t}$ is a vector that contains all the significant country-specific factors ($Z_{i,t}$) and common factors (F_t) identified in the two-step procedure detailed in the previous section. Our aim is to find a sample from the non-crisis period that has characteristics $X_{i,t}$ that most closely match the characteristics of the observations in the crisis sample. Once we obtain the coefficients of interest and the predicted probabilities of the cumulative standard logistic distribution $\Pr(\mathrm{D}_{\mathrm{GFC},t} = 1)$) from equation (4.7), we can compute the fitted cumulative probability that the observation is not in the crisis "treatment":

$$p_{i,t} = 1 - \Pr(\mathrm{D}_{\mathrm{GFC},t} = 1 | X_{i,t}) \quad (4.8)$$

Once we have estimated these probability values for all the countries at each point in time, we implement a nearest neighbout matching procedure that we will describe in the following paragraphs. The results of the logit regression for the whole period of the GFC using the significant coefficients obtained from the stepwise procedure in Table 4.3 are summarised in Table 4.4.

Table 4.4: Logistic regression estimates for calculating propensity scores

This table reports the results from panel logit regressions of the Global Financial Crisis indicator variable on a set of the significant explanatory variables obtained in the two-step procedure detailed in Table 4.3 and Section 2.4.2. The regression equation is $\Pr(D_{GFC,t} = 1 | X_{i,t}) = (1 + \exp(-\alpha_o - \beta_{i,t} X_{i,t}))^{-1}$ where $D_{GFC,t} =$ the crisis indicator for the global financial crisis ($D_{GFC,t}$); it takes the value of 1 for the period between July 2007 and May 2012 and 0 otherwise. ***1%; **5%; and *10% significance. The propensity scores are obtained by calculating the predicted values of the regression and equation (4.8).

Dependent variable $D_{GFC,t}$	**All countries**	**Developed**	**Emerging**	**EUR**	**USD**	**LOCAL**	**Local developed**	**Local emerging**	**PIIGS**
Spread (t-1)	0.2216***	0.3595***	0.1140***	0.2651***	0.0279	0.3146***	0.3991***	0.3515***	0.0138
	(0.0129)	(0.0282)	(0.0171)	(0.0403)	(0.0223)	(0.0196)	(0.0369)	(0.0319)	(0.0367)
Country-specific factors									
Exchange rate	-0.0174**		-0.0088		-0.0244	-0.0233*		0.0114	-0.0216
	(0.0080)		(0.0110)		(0.0156)	(0.0127)		(0.0175)	(0.0414)
Debt/GDP	-0.0067***	0.0013					0.0147***		
	(0.0012)	(0.0018)					(0.0024)		
Investment/GDP		0.0353***							
		(0.0030)							
Imports/GDP		-0.0021							
		(0.0068)							
Debt/exports	0.0024**	-0.0003	-0.0046***		0.0011		-0.0110***		
	(0.0010)	(0.0016)	(0.0012)		(0.0014)		(0.0023)		
GDP per capita growth	-0.0949***	-0.1651***		-0.1904***		-0.1130***	-0.1074***	-0.0841***	-0.3719***
	(0.0057)	(0.0120)		(0.0171)		(0.0094)	(0.0140)	(0.0137)	(0.0382)
Current account/GDP			-0.0074						-0.2245***
			(0.0068)						(0.0247)

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
Reserves		0.0132***				0.0181***		0.0512***	
		(0.0024)				(0.0016)		(0.0052)	
Local equity premium	-0.0806***	-0.0433***	-0.0343***		-0.0332***	-0.2134***	-0.2013***	-0.0347***	-0.0655
	(0.0080)	(0.0070)	(0.0068)		(0.0096)	(0.0243)	(0.0295)	(0.0091)	(0.0689)

	All countries	Developed	Emerging	EUR	USD	LOCAL	Local developed	Local emerging	PIIGS
Common factors									
Global equity premium	-0.0442***					-0.1895***	-0.1712***		0.0554
	(0.0093)					(0.0257)	(0.0312)		(0.0787)
European bond index	-0.0828***	-0.6796***	-0.0441		-0.0431	-0.0832**		-0.0424	-0.7602***
	(0.0276)	(0.0809)	(0.0281)		(0.0431)	(0.0384)		(0.0399)	(0.2014)
US equity premium			-0.0091		-0.0212	0.1585***	0.1722***		0.0158
			(0.0139)		(0.0213)	(0.0263)	(0.0328)		(0.0709)
Regional bond portfolio	0.2365	6.6870***		0.8082**		0.3729	0.3643		8.5107***
	(0.2400)	(0.8151)		(0.3677)		(0.3269)	(0.3578)		(2.1493)
Global risk aversion	-0.0068***		-0.0006	0.0068**	-0.0016	-0.0046			
	(0.0021)		(0.0031)	(0.0030)	(0.0047)	(0.0030)			
Pseudo R2	0.0829	0.1439	0.0262	0.1338	0.0143	0.1137	0.1069	0.1145	0.3785
Number of observations	6235	3646	2679	1869	1126	3330	2058	1432	788
Obs with dependant=0	3818	2289	1529	1161	654	2003	1291	803	493
Obs with dependant=1	2417	1357	1150	708	472	1327	767	629	295

Our procedure for testing differences in spreads is based on the average treatment effect on the treated (ATET) framework. This procedure uses the probabilities obtained in equation (4.8) and the original sovereign spread values to make a selection of counterfactual values based on propensity score matching. This procedure has certain advantages over traditional sampling or predicted values difference testing since it effectively addresses the problem of selection bias of comparable sample groups during the non-crisis period. One key advantage of this method is that we can compare the actual value of the spreads without forgoing the theoretical richness contained in the observable characteristics of a pricing factor model. Finally, with ATET it is possible to determine exactly which observations in the non-crisis are more closely related in terms of common determinants to those in the crisis periods, which can have important implications regarding policy making or early warning systems.

This method was originally developed by (Rosenbaum and Rubin, 1983) in order to address the non-randomness of treated vs. non-treated groups in medical trials, and since then has been applied to other areas of the social sciences such as labour economics and finance. In this paper, we modify the framework proposed by (Nssah, 2006) on how to apply ATET to economic policy programs and reframe it for contagion testing. In the context of corporate bond markets, this framework has been used to test the impact of credit supply shocks in the capital structure of the firm (Almeida et al., 2009).

Here, the "treated" group is characterised by a dummy that represents the crisis dates (D=1) and the "non-treated" are represented by the non-crisis dates (D=0). Therefore, by dividing the spreads ($spreads_{i,t}$) from equation (4.8) into two vectors that represent the crisis period ($\{spreads_{crisis}\}$) and non-crisis period ($\{spread_{noncrisis}\}$) using the algorithm in equation (4.12) we have:

$$\{g_i\} = (\{spreads_{crisis}\} - \{spread_{noncrisis}\}) \tag{4.9}$$

where the average value of the vector $\{g_i\}$ is equal to the ATET. Additionally, if we assume that there is unit homogeneity,[36] since in a global crisis countries do not have the freedom to "choose" whether to participate in it or not, we can rewrite $\{g_i\}$ in conditional probability form where:

$$ATET = E\left(\{g_i\}\middle|X, D=1\right) = E\left(\{spreads_{crisis}\}\middle|X, D=1\right) - E\left(\{spreads_{noncrisis}\}\middle|X, D=0\right) \tag{4.10}$$

where X is the vector of common observable characteristics represented by the explanatory variables from equation (4.8) and the averages of $E\left(\{spreads_{crisis}\}\middle|X, D=1\right)$ and $E\left(\{spreads_{noncrisis}\}\middle|X, D=0\right)$ represent respectively the mean of the "treated" and the counterfactual mean of the "non-treated" or, in our setup, the crisis and non-crisis period. ATET using propensity-matching estimators represents an interesting framework for testing contagion because the method yields strong estimates under the assumption of conditional independence (Abadie et al., 2004). The assumption can be formally defined as:

$$\left(spreads_{crisis}, spreads_{noncrisis}\right) \perp D\middle|X\right) \tag{4.11}$$

In other words, conditional on observable characteristics (X), participation (D) is independent of the potential outcomes of $\left(spreads_{crisis}, spreads_{noncrisis}\right)$. In order to be coherent with the principle of conditional independence, the basic idea behind

[36] Unit homogeneity refers to the fact that participants cannot choose to participate in the experiment, so the experimental group is comprised of both volunteers and non-volunteers; there is no bias based on the willingness of the participants to be a part of a given experiment.

propensity matching is to randomly select a sample from the non-crisis (non-treated) period that most closely resembles the characteristics of our sample in the crisis (treated) period conditional on the common factors, reducing selection bias.

Using the probability values from equation (4.8) we can implement the algorithm in equation (4.12) for finding the vector with nearest neighbour matching estimators (NNB):

$$c(p_{\text{matched},t}) = \left\{ j \middle| \min \left\| p_{\text{crisis},t} - p_{\text{noncrisis},t} \right\| \right\} \qquad (4.12)$$

Where $c(p_{\text{matched},t})$ represents the vector of matched crisis and non-crisis spreads based on the nearest difference propensity scores which are simply one minus the cumulative probabilities obtained using equation (4.7), where ($P_{crisis} = 1 - \Pr(\mathrm{D}_{\text{GFC},t} = 1 | X_{i,t})$) are the cumulative probabilities for those observations in the crisis period and ($P_{noncrisis} = 1 - \Pr(\mathrm{D}_{\text{NOGFC},t} = 0 | X_{i,t})$) are those of the non-crisis period. The vector that represents the non-crisis period ($\{spread_{noncrisis}\}$) is constructed by selecting the spreads at the corresponding dates of the $p_{\text{noncrisis}}$ cumulative probabilities obtained with equation (4.10). Therefore, we can find evidence if there is difference in spreads by testing if the average of the matched vector $\{g_i\}$ is statistically significant via a simple ANOVA test where the null of no differences in spreads versus the alternative is formally defined as:

$$H_0 : \overline{spreads}_{crisis} = \overline{spreads}_{noncrisis}$$
$$H_1 : \overline{spreads}_{crisis} \neq \overline{spreads}_{noncrisis} \qquad (3.13)$$

In this hypothesis, $\overline{spreads}_{crisis}$ and $\overline{spreads}_{noncrisis}$ are the mean values of the observations in vectors $\{spreads_{crisis}\}$ and $\{spreads_{noncrisis}\}$ according to the matched propensity scores in vector $c(p_{\text{matched},t})$. In this way we observe the impact of the spreads in the crisis periods relative to the observations that most closely resemble the crisis characteristics in the non-crisis periods as well as overlapping periods. Furthermore, we observe the effect of changing averages in spreads during three different phases of the GFC (subprime, credit, and ESD) as well as for the whole period of the GFC between July 26, 2007, until May 17, 2012. We compare the ATET results with those obtained by using other criteria for setting the samples of the crisis and non-crisis period: first that allow unequal crisis and non-crisis samples; and second that force the crisis and non-crisis samples to be equal, as is often the case for other contagion-testing methods that use correlations or that test for increases in factor loadings (Dungey et al., 2005).

4.4 Results

Table 4.5, Panel A, reports the results obtained from the matching procedure using country-specific determinants. In the case of the total period (GFC) all counterfactuals are drawn solely from the non-crisis period. In the case of the crisis phases we allow for counterfactuals from the non-crisis periods and other phases in order to see if there are significant differences among crisis periods. Our results show that for our portfolios there was a significant difference in spreads for the whole period of the GFC with the exception of local-currency-issued emerging-market debt, in which the difference with comparable counterfactuals is statistically insignificant. The significant changes in spreads in most groups can be explained by cross-market linkages through fundamentals related to liquidity or “wake-up” contagion.

Although the case of local-currency emerging-market debt could be viewed as counterintuitive because this kind of debt has been traditionally considered to be a high-risk investment, it is important to recall that we are comparing the *characteristics* of a certain crisis period with the *characteristics* of a different period that *most closely* resembles those of the crisis period in the past. In the case of local-currency emerging-market debt, this means that there were periods with *similar severity in terms of variance* which do not necessarily translate to *low* or *high magnitude* in the changes of spreads. The average spread in the GFC for the local-currency emerging-market debt issuers was 335.01 basis points, and 316.46 for those selected in the non-crisis period using neighbour matching estimators (NNB).

Table 4.5: Average effect on sovereign spreads on portfolios

Panel A reports the results from matching the inverse cumulative probabilities (propensity scores), which are obtained by applying equation (4.8) to the predicted probabilities obtained from the logit regressions for each of the (9) portfolios during the whole GFC period. For the matching procedure between crisis and non-crisis periods we use the algorithm for the nearest neighbour matching estimator (NNB) in equation (12). The results reported in the NNB column correspond to the average monthly spread of the counterfactual vector obtained using equation (12). The subprime, credit, and ESD columns correspond to the average monthly spread for each country during the different crisis dates. The average treatment effect on the treated (ATET) is simply the arithmetic difference between a crisis period average and its corresponding NNB average. The statistical significance of the ATET is tested using an ANOVA test for equality of means of the vector containing the observations of crisis periods and their respective vector of NNB counterfactuals. Panel B reports the results for the statistical difference in the average spreads using an unequal sample for the non-crisis period, which is dated from January 2000 to June 2007, versus the average spread for each portfolio in the respective crisis period. Panel C reports the results for the statistical difference in the average spreads using an equal sample for the non-crisis period versus the average spread for each portfolio in the respective crisis period, so in the case of the case of the whole GFC, the crisis period is 59 months, the non-crisis period is exactly 59 months from August 2002 to June 2008. Panel D reports the results for the statistical difference in the average spreads using an equal sample for the non-crisis period versus the average spread for each portfolio in the respective crisis period but using the data from a previous crisis phase (credit crisis is compared with data of equal size with the subprime phase and the ESD takes data from the subprime and credit phases respectively). ***1%; **5%; and *10% significance

Panel A: NNB matching noncrisis period

Country	Subprime	NNB	ATET	Credit	NNB	ATET	ESD	NNB	ATET	GFC	NNB	ATET
All	28.18	17.53	10.65	23.87	-5.34	29.21	68.42	2.31	66.10***	49.05	-21.73	70.78***
Developed	-6.34	-52.08	45.74***	-21.57	-29.16	7.58	27.95	-4.34	32.29**	8.90	-62.05	70.95***
Emerging	275.97	312.07	-36.10	350.04	267.36	82.68***	358.86	275.67	83.19***	337.25	254.39	82.86***
Euro	9.04	-72.30	81.35***	-5.13	34.54	-39.66*	71.23	3.42	67.81***	39.65	-64.33	103.98***
Local	31.45	-1.18	32.62	20.09	-11.78	31.87	51.64	21.85	29.79***	39.89	-28.15	68.05***
USD	280.46	344.73	-64.27**	563.91	284.89	279.02***	326.35	340.27	-13.92	367.80	218.31	149.49***
Local developed	-22.07	-59.68	37.61*	-37.45	-59.74	22.29	-17.73	-30.12	12.39	-23.11	-68.94	45.83***
Local emerging	282.11	276.79	5.32	289.60	279.63	9.97	376.60	326.29	50.31***	335.01	316.46	18.54
PIIGS	26.53	109.06	-82.53	27.36	169.10	-141.74**	242.03	50.09	191.94***	143.60	3.40	140.20***

Panel B: Unequal sample non-crisis period

Country	Subprime	Unequal	Difference	Credit	Unequal	Difference	ESD	Unequal	Difference	GFC	Unequal	Difference
All	28.18	-50.00	78.18***	23.87	-50.00	73.87***	68.42	-50.00	118.41***	49.05	-50.00	99.05***
Developed	-6.34	-87.67	81.33***	-21.57	-87.67	66.10***	27.95	-87.67	115.62***	8.90	-87.67	96.57***
Emerging	275.97	220.42	55.55***	350.04	220.42	129.62***	358.86	220.42	138.44***	337.25	220.42	116.83***
Euro	9.04	-68.89	77.93***	-5.13	-68.89	63.76***	71.23	-68.89	140.12***	39.65	-68.89	108.54***
Local	31.44	-50.24	81.68***	20.09	-50.24	70.32***	51.64	-50.24	101.87***	39.89	-50.24	90.13***
USD	280.46	261.59	18.86	563.91	261.59	302.32***	326.35	261.59	64.76***	367.80	261.59	106.21***
Local developed	-22.07	-107.14	85.07***	-37.45	-107.14	69.69***	-17.73	-107.14	89.40***	-23.11	-107.14	84.03***
Local emerging	282.11	216.30	65.81***	289.60	216.30	73.30***	376.60	216.30	160.29***	335.01	216.30	118.71***
PIIGS	26.53	-63.15	89.68***	27.36	-63.15	90.51***	242.03	-63.15	305.18***	143.60	-63.15	206.74***

Panel C: Equal sample non-crisis period

Country	Subprime	Equal	Difference	Credit	Equal	Difference	ESD	Equal	Difference	GFC	Equal	Difference
All	28.18	-68.53	96.71***	23.87	-65.29	89.16***	68.42	-77.31	145.73***	49.05	-49.18	98.23***
Developed	-6.34	-101.06	94.72***	-21.57	-97.85	76.28***	27.95	-111.19	139.14***	8.90	-81.98	90.89***
Emerging	275.97	164.96	111.01***	350.04	168.42	181.62***	358.86	165.87	192.99***	337.25	186.28	150.97***
Euro	9.04	-90.92	99.96***	-5.13	-87.62	82.49***	71.23	-107.51	178.74***	39.65	-76.99	116.64***
Local	31.44	-62.68	94.12***	20.09	-59.31	79.39***	51.64	-65.93	117.57***	39.89	-41.01	80.90***
USD	280.46	189.74	90.71***	563.91	189.47	374.44***	326.35	207.92	118.43***	367.80	255.21	112.59***
Local developed	-22.07	-111.37	89.30***	-37.45	-108.28	70.83***	-17.73	-114.54	96.81***	-23.11	-86.37	63.27***
Local emerging	282.11	165.43	116.69***	289.60	170.10	119.50***	376.60	161.80	214.80***	335.01	171.49	163.52***
PIIGS	26.53	-81.03	107.56***	27.36	-77.39	104.75***	242.03	-100.67	342.70***	143.60	-70.72	214.31***

Panel D: Overlapping sample with previous crises periods (credit and ESD)

Country	Credit	Overlapping	Difference	ESD	Overlapping	Difference
All	23.87	33.35	-9.48	68.42	14.91	53.50***
Developed	-21.57	-1.70	-19.87	27.95	-23.21	51.16***
Emerging	350.04	284.98	65.06**	358.86	288.52	70.34***
Euro	-5.13	13.58	-18.71	71.23	-7.65	78.88***
Local	20.09	37.02	-16.93	51.64	15.16	36.48***
USD	563.91	288.64	275.26***	326.35	377.48	-51.13*
Local developed	-37.45	-17.28	-20.17	-17.73	-38.83	21.09***
Local emerging	289.60	291.34	-1.74	376.60	268.02	108.57***
PIIGS	27.36	31.85	-4.49	242.03	14.56	227.47***

In the case of the GFC (see Table 4.5, Panel A) the most significant statistical difference in spreads was from the troubled countries (PIIGS) and those countries that issue USD-denominated debt. In the case of the PIIGS the channel of contagion was related mainly to leading macroeconomic indicators of liquidity (current account and investment to GDP) and changes in the sovereign spreads in the other troubled countries that are part of the PIIGS portfolio. On the other hand, the USD-denominated debt change in spreads is attributable to cross-market linkages among fundamentals related to liquidity and evidence of a latent factor with a "positive contagion" effect. In the case of the subprime phase the change in spreads was statistically significant in developed and Eurozone countries. In the case of the USD-denominated debt portfolio we observe a significant reduction of spreads of -64.27 basis points. One reason could be investors replacing US-backed mortgage securities with other USD-denominated debt.

In the credit crisis period the change in spreads was statistically significant in emerging countries, and especially USD-denominated debt issues, probably because liquidity in the US market dried up after the Lehman collapse. We observe an average increase in the cost of USD debt issues of 279.02 basis points. Curiously, the PIIGS and the Eurozone portfolio reported a reduction of -141.74 and -39.66 basis points respectively during that phase. This reduction could be attributable to rebalancing effects by investors from the US to the Eurozone amid the drain of liquidity in the US markets. Finally, we can observe that in the ESD phase all the portfolios reported significant changes, with the exception of the USD and local-currency debt issues in developed countries. As expected, the PIIGS portfolio reported the biggest increase in spreads with a total of 191.94 basis points during the ESD phase.

The results reported in Table 4.5, Panel A, using the ATET framework are strikingly different to those reported in Panel B and Panel C, in which we allow for both equal and unequal samples in all the phases using standard exogenous dating methods. The only exception is the USD-denominated debt portfolio that does not report significant statistical changes during the subprime phase. Even when we allow for overlapping samples, the results are grossly underestimated or overestimated when compared with results obtained with matching estimators. The most compelling argument for the use of matching estimators is that we can have a reliable measure of the economic impact in the spreads that effectively incorporates the information of the determinants in the reported change. Therefore, in Table 4.6 we show the results of robustness checks using alternative matching kernels that impose a region of common support, which means that we limit our draws of counterfactuals to those observations that are between the minimum and maximum probabilities of the crisis period values, as outlined in Appendix D. These alternative kernels further refine the sample universe by setting minimums and maximums.

Table 4.6: Robustness checks for ATET on portfolios

This table reports the results from matching the inverse cumulative probabilities (propensity scores) which are obtained by applying equation (4.8) to the predicted probabilities obtained from the logit regressions for each of the (9) portfolios during the whole GFC period. The results reported in the NNB, Gauss, and EPNK columns correspond to the average treatment effect on the treated (ATET), which is the difference between the crisis period average monthly spread vectors and the counterfactual vectors for each country. These results are obtained using equation (4.12) in the case of the NNB and equations (4.14) and (4.15) in Appendix D when using Gaussian (GAUSS) and Epanechnikov kernels as outlined in Section 3.5. The statistical significance of the ATET is tested using an ANOVA test for equality of means of the vector containing the observations of crisis periods and their respective vector of counterfactuals using three matching methods (NNB, GAUSS, and EPNK). For reasons of space we do not report the average of each counterfactual vector as shown in Table 4.5, just the final ATET result for each method. ***1%; **5%; and *10% significance.

	Subprime			**Credit**			**ESD**			**GFC**		
Country	**NNB**	**GAUSS**	**EPNK**	**NNB**	**GAUSS**	**EPNK**	**NNB**	**GAUSS**	**EPNK**	**NNB**	**GAUSS**	**EPNK**
All	10.65	35.15**	33.08**	29.21	31.81*	39.95**	66.10***	63.96***	72.13***	70.78***	73.43***	81.57***
Developed	45.74***	39.55	31.96**	7.58	8.85	7.40	32.29**	40.42***	39.88**	70.95***	65.43***	66.23***
Emerging	-36.10	-5.85	-6.68	82.68***	80.35***	73.34***	83.19***	92.63***	92.24***	82.86***	90.22***	93.80***
Euro	81.35***	69.81	68.37***	-39.66*	-50.01*	-47.61	67.81***	64.67***	86.08**	103.98***	98.15***	51.25***
Local	32.62	38.85	44.25***	31.87	22.54	28.38*	29.79***	39.52***	40.20***	68.05***	69.76***	65.87***
USD	-64.27**	-50.47	-52.10***	279.02***	281.64***	270.52***	-13.92	20.61**	12.31	149.49***	142.52***	154.17***
Local developed	37.61*	30.35	26.26*	22.29	25.33*	23.57	12.39	11.81*	9.12	45.83***	47.78***	40.56***
Local emerging	5.32	2.91	0.93	9.97	-7.80	2.68	50.31***	48.60***	30.73***	18.54	14.32	-5.55*
PIIGS	-82.53	-24.05	-24.79***	-141.74**	-163.04***	-144.89***	191.94***	198.42***	251.98***	140.20***	159.67***	110.18***

Where two of the three methods yield similar results, we categorise the result as being a meaningful difference, otherwise we keep the original results. From the robust specification, we can observe that in the case of local-currency emerging-market debt the difference in spreads is statistically insignificant and in the order of 0.93 basis point during the subprime crisis as opposed to 5.32 basis points using the conventional methods. In the subprime and credit phases there is evidence of a significant increase in the spread difference for the all-countries portfolio using the robust method. The results concerning the behaviour of local-currency emerging-market debt are in line with other studies that suggested the possible decoupling of emerging markets during the early phases of the GFC (Dooley and Hutchison, 2009) and to a more recent study by Dungey et al. (2010), which found that the US subprime crisis had only a small impact in the volatility of emerging-market sovereign bond markets, albeit the studies employed totally different methodologies.

4.5 Conclusions

By using a factor model based on country-specific and common market determinants of sovereign spreads in the context of propensity-matching estimators we propose a novel framework to test for differences in spreads. Our findings suggest that the most common country-specific factor among portfolio groups is the local equity premium, with the exception of the Eurozone and PIIGS, where the changes in neighbouring countries (regional portfolio) have a larger effect. However, in the specific case of the GFC, there were different channels for transmission of contagion. The most common channel of contagion transmission among most portfolio groups was macroeconomic fundamentals related to liquidity or “wake-up” contagion, where investors pay close attention to the country’s ability to meet its financial obligations. There was evidence of a latent-factors or "pure" contagion in all of the portfolios with the exception of the all-countries portfolio, suggesting that most of the latent factors in the remaining portfolios can be explained by changes in the global bond portfolio. The channel of contagion in the case of local currency issuers during all the phases was related to common factor contagion (global equity premium, the European bond index, regional portfolio, and changes in the perception of global risk aversion). In summary, our findings show that when we divide the crisis into sub-periods we can observe that the first two periods (Subprime and Credit) were driven by country specific fundamentals "wake up" contagion and that the last period by "pure" or latent contagion that can be explained by risk aversion in the global bond portfolio. Additionally, in our proposed framework, we define our test for differences in spreads as a statistically significant change in the average change in spreads between the observations in the counterfactual non-crisis period and those of the crisis period. We do this in order to determine the actual economic significance in basis points of the different phases of the crisis versus non-crisis periods. In order to do this we define this average change as our average treatment effect on the treated (ATET), where the "treatment" is the crisis period. In this way, we are able to obtain estimates based on a similar distribution between the crisis and counterfactual group and reduce the problem of selection bias inherent in ANOVA testing. We test the robustness of the results using equal, unequal, and overlapping non-crisis periods and also using different kernel specification of matching estimators without significant changes in our main results.

In summary, the evidence shows that the most meaningful periods for differences in sovereign spreads was the ESD phase, in which the spreads rose substantially from the previous crisis periods. The exception was the USD-denominated debt portfolio, which displayed a major variation in spread during the credit crisis phase. The portfolios that were most affected in terms of spreads during the GFC were the troubled countries (PIIGS) and the USD-denominated debt portfolio. For the market-weighted portfolio of the 43 countries in our sample the average spread rose by 70.78 basis points compared with similar events during the non-crisis period. Finally, we found evidence that the portfolio of local-currency emerging-market debt did not exhibit any significant difference in spreads during the GFC as a whole, even under robust specifications. This means that based on the common characteristics of the counterfactuals, the emerging markets that issued debt in local currency have dealt with similar economic conditions in the past.

Chapter 5 The effects of the GFC on Colombian local-currency bonds prices: an event study

5.1 Introduction

The impact of financial contagion during the Global Financial Crisis (GFC) has produced a vast body of literature. One strand of the literature has dealt with the effect of the GFC on emerging markets. Of particular interest is the fact that stock markets in emerging economies seemed to be decoupled or insulated from the events that followed the Lehman bankruptcy and subsequent credit crunch phase of the financial crisis (Dooley and Hutchison, 2009).

While some studies have corroborated these findings using different methods in different emerging markets (Fidrmuc and Korhonen, 2010; Köksal and Orhan, 2013; M Ayhan Kose, 2011; Kose et al., 2012; Samarakoon, 2011), other studies have found no evidence of decoupling of emerging markets during the GFC and have concluded that financial contagion was prevalent, especially in Latin America (Beirne and Gieck, 2014; Dufrénot et al., 2011; Felices and Wieladek, 2012; Levy Yeyati and Williams, 2012). Most of these studies rely on factor models that contain local and global factors and are usually based on daily, monthly, or quarterly stock and bond prices depending on the level of aggregation of the macroeconomic data that supports the factor model in which the study is based.[37]

However, there are not many studies that test for contagion in emerging markets using intraday stock or bond data, or that use the event study approach. The main advantage of using intraday data is that we can obtain more precise estimates of the real impact of news on asset prices. By using intraday data, one can control for

[37] This is the case for studies that try to identify transmission channels of contagion through macroeconomic indicators such as trade and other GDP-based variables in the factor model. Some examples can be found in Bekaert et al. (2011) and Csontó (2014).

the confounding effects of other similar announcements during aggregation periods, and therefore reduce the "noise" attributable to events other than the one under study (Bomfim, 2003; Chuliá et al., 2010; Fair, 2002). In the case of the GFC these confounding effects are present during the different events of the crisis. For example, there were days such as 15 September of 2008 (see Table 5.3) with more than one announcement during the same day.

We utilise an event study approach using high-frequency data on peso-denominated Colombian government bonds to measure the effects of news during the GFC. The main objective of our study is to add to understanding of the coupling and decoupling hypothesis in emerging markets during the GFC. The second objective is to add to understanding of the effects of macroeconomic news on Colombian bond prices.

The Colombian peso-denominated government bond market makes an interesting case study because it is third in market capitalisation in Latin America just behind Brazil and Mexico. In August 2014, the market value of local-currency debt of the five principal issuers in Latin America amounted to a total US$381.88 billion, of which Colombian bonds represented US$75.11 billion (Citibank, 2014).

Few studies have analysed the effects of the GFC-sourced contagion from both an event study and microstructure perspective. Aizenman et al. (2012) analysed the contagious effects of news derived from the Euro crisis on the prices of bonds of developed and developing markets and found evidence that news from the Euro crisis had a moderate effect on developing-country bond prices. Beetsma et al. (2013) examined the asymmetric effects of good and bad news derived from the Euro crisis in the Eurozone and found that bad news was the significant driver behind increases in domestic interest rates in troubled countries. Both of these studies used daily data and relied mainly on indices to measure the effects of the events on yield and prices. In our event study using individual bonds we find that, in most cases, bad news related to the crisis had a positive impact in Colombian bond prices during the GFC.

Our study has some commonalities with a recent study of the Italian goverment bond market by Pelizzon et al. (2013), using tick data from June 2011 to December 2012 to measure the effects of the Euro crisis on liquidity and order flow. They find that a spike in credit risk caused market dealers to reduce their provision of liquidity, exacerbating the drop in Italian bond prices. Their study was focused mainly on macroeconomic news events pertaining to policy action during a short episode of the Euro crisis. Our study aims to incorporate some elements of these previous studies while also controlling for the confounding effects of surprises in normal macroeconomic announcements. Even after controlling for confounding effects, we find evidence of positive abnormal returns for Colombian bonds during the GFC.

The remainder of the paper is divided as follows. Section 5.2 describes our dataset and the particularities of the Colombian government-bond market. Section 5.3 describes the method employed in the event study of news surrounding the GFC and other macroeconomic announcements. Section 5.4 presents our results and Section 5.5 concludes.

5.2 Data[38]

Bond data in Colombian pesos is from the Colombian interdealer electronic negotiation system (SEN in Spanish). The administrator of the system is the

[38] All the information depicted in this section is taken from rules for operation of the national electronic negotiation system (SEN) in Spanish and can be found at http://www.amvcolombia.org.co/attachments/data/20110214135937.pdf (AMV, 2008)

Colombian Central Bank, which can also act as a market dealer. The SEN is a two-tier market platform. The first tier is limited to the market dealers responsible for providing liquidity to the bond market. In the second tier are all other agents that are approved as market agents but not as market dealers; only the market dealers are able to put bids and ask quotes in both the first and second tier. In order to be a market dealer in Colombia, financial institutions have to meet a set of minimum requirements in terms of capital and liquidity, as well as mandatory participation in the Dutch auctions of Colombian peso government bonds in the primary market. Table 5.1 reports the consolidated yearly turnover of the SEN market in Colombian pesos: the average monthly exchange rate as at August 2014 was USD/COP 1,899.07, making the total traded dollar value for the 2014 year to date approximately US$559 billion.

Table 5.1: Historical COP turnover value SEN interdealer local-bond currency market, 1999 to August 2014

Year	Number of transactions		Total turnover nominal value (Trillion COP)		Total turnover market value (Trillion COP)	
	Total	Monthly average	Total	Monthly average	Total	Monthly average
1999	5.247	437	5.7	0.5	6.4	0.5
2000	12.076	1.006	14.6	1.2	15.6	1.3
2001	53.877	4.490	84.4	7	87.7	7.3
2002	87.040	7.253	155.5	13	167	13.9
2003	142.830	11.903	259.8	21.6	275.1	22.9
2004	309.710	25.809	852.1	71	898.8	74.9
2005	453.612	37.801	1,071.90	89.3	1,203.60	100.3
2006	277.480	23.123	718.1	59.8	783.7	65.3
2007	116.948	9.746	433.95	36.2	458.6	38.2
2008	115.141	9.525	608.47	50.7	614.7	51.2
2009	371.836	30.986	1,913.75	159.5	2,089.60	174.1
2010	219.957	18.330	1,264.82	105.4	1,367.60	114
2011	150.769	12.564	883.05	73.6	838.6	69.9
2012	254.675	21.223	1,329.65	110.8	1,468.50	122.4
2013	163.612	13.634	1,000.01	83.3	1,069.10	89.1
2014(Aug)	187.192	23.399	1,063,46	132,9	1,111,70	139

Source: Banco de la Republica (2014)

In the first tier there are two principal types of transactions: buy and sell operations with same-day clearance (T+0) and simultaneous operations with clearance periods greater than one day (T>=0). For buy and sell operations the participants must make offers of a minimum one billion pesos (approximately US$500,000) in increments of 500 million pesos (US$250,000). The same rules apply for simultaneous operations (repo operations), except the investor who receives the bonds as collateral can trade with them until the repurchase takes place. All quotations are in clean prices and for the total minimum nominal value of the dematerialised titles that make up a specific issue (which in Colombia is 500 million pesos or approximately US$250,000). All the operations are blind, in the sense that no other participants can see the identity of the involved party: the identity of buyer and seller is disclosed when a trade takes place but only to the parties involved in the operation. The only valid types of orders are: FOK (fill or kill), in which a trade is taken partially and the remainder of the total offered quantity disappears from the screen; GTS (good until specific), in which the interested party puts the quote for a specified amount of time and the remainder of the total quantity is left if just a minimum is taken; and GTC (good until cancelled), which is the same as GTS but only for the duration of the trading day, and at the close of the system the order is cancelled.

In the case of principal, coupons, and very short-term bonds, the quotations are made in yields instead of clean prices. For the purpose of this study all transactions are buy and sell transactions using the settlement clean price for each trade, and we exclude all the instruments that are traded using yields, which in the Colombian case are zero coupon bonds and coupon strips created synthetically from regular government bonds and that rarely trade in first-tier sessions. All the data for this study was aggregated hourly using the continuously compounded returns of settlement trade prices during a specific hour. Our sample period runs from January 2007 until December 2013 and uses both active and expired issues during the window of observation to avoid possible survivorship bias. Also we address liquidity issues by selecting bonds that have more transactions than calendar days from January 1, 2007, until December 31, 2013. In the case of those bonds issued before January 1, 2007, and that settled before December 31, 2013, we apply the same rule but count the days from the beginning of our window of observation until the date of settlement. Similarly for bonds issued after January 1, 2007, we apply the same rule but count the days from the date of issue until December 31, 2013. In Colombia all

Colombian peso government bonds pay the coupon once a year on the date specified in the issue, so for example in the ticker TFIT11241018 the "TFIT" means that it is a bond with annual coupons, and the last six digits specify the date of payments of the coupons and principal. (In this example the coupon is payable every October 24 with principal payable in the year 2018.) In Table 5.2 we observe the remaining issues and their descriptive statistics (we do not report the mean returns, which in all cases are close to zero to the fourth decimal, and all the returns are expressed in hourly compounded returns). From the descriptive statistics we can observe that in almost all cases there is negative skewness, which is an indicator that negative returns have a higher impact than positive returns, and compounded hourly returns range between -7.43% and 4.25%.

Table 5.2: Descriptive statistics for selected issues of Colombian peso government bonds

			Hourly returns					
Issue	# Hourly observations	Per cent of observations	Standard deviation	Minimum	Maximum	Skewness	Kurtosis	Coupon
TFIT03140509	2096	4.12%	0.04%	-1.19%	0.23%	-10.81	294.59	8.75%
TFIT04150812	3794	7.46%	0.08%	-0.71%	0.80%	0.18	23.79	9.25%
TFIT04170413	2755	5.41%	0.07%	-1.48%	0.95%	-2.87	96.29	6.00%
TFIT04180511	3143	6.18%	0.08%	-1.65%	0.73%	-2.54	70.28	11.00%
TFIT05100709	1231	2.42%	0.06%	-0.49%	0.38%	-1.05	12.30	12.50%
TFIT05241110	3719	7.31%	0.08%	-0.82%	0.64%	-0.62	17.93	7.50%
TFIT06140514	5492	10.79%	0.10%	-1.28%	1.06%	-0.63	28.42	9.25%
TFIT06141113	3220	6.33%	0.19%	-2.87%	1.94%	-1.26	31.76	10.25%
TFIT07150616	4104	8.07%	0.17%	-7.43%	1.31%	-21.10	941.21	7.25%
TFIT07220808	1370	2.69%	0.03%	-0.20%	0.17%	-1.08	10.79	15.00%
TFIT10040522	1575	3.10%	0.28%	-2.60%	1.76%	-1.06	18.82	7.00%
TFIT10281015	4098	8.05%	0.25%	-3.73%	3.69%	0.40	49.56	8.00%
TFIT11241018	2245	4.41%	0.37%	-6.61%	4.25%	-3.26	77.61	11.25%
TFIT15240720	6296	12.37%	0.29%	-2.90%	4.11%	-0.27	22.79	11.00%
TFIT15260826	1061	2.09%	0.41%	-5.96%	2.74%	-3.54	58.72	7.50%
TFIT16240724	4686	9.21%	0.28%	-3.65%	2.57%	-0.91	19.32	10.00%

Source: Data retrieved from the Colombian Central Bank (Banrep, 2014). In Colombia all Colombian peso government bonds pay the coupon once a year on the date specified in the issue, so for example in the ticker TFIT11241018 the "TFIT" means that it is a bond with annual coupons, and the last six digits specify the date of payments of the coupons and principal.

For a US market proxy we choose the iShares Core S&P 500 ETF (ticker: IVV), which is the exchange-traded fund that tracks the S&P 500. Compared with traditional indexes, IVV has the distinct advantage of being tradable and available at high frequency from Thomson Reuters tick history database. This is important since we assume that the US stock market was the principal source of transmission until the end of the GFC. We exclude the European sovereign debt crisis episode due to the lack of a high-frequency traded index proxy that accurately represents the transmission mechanism from European debt markets to the rest of the world. All the times reported in this study are in Eastern Standard Time, adjusted for daylight savings time as appropriate.

To identify the events of the GFC we use the market events outlined in the financial turmoil timeline of the Federal Reserve of New York (FEDNewYork, 2011) with the exception of the first event,[39] which is taken from the crisis timeline of the Federal Reserve of St. Louis (FEDSt.Louis, 2009). We adopt a GFC crisis period consistent with the timeline, which is from July 24, 2007, until December 28, 2010. The timestamp (see Table 4.3) for the exact release of each event was retrieved from Bloomberg at the first time the event appeared in Bloomberg and discarding subsequent follow-ups. Finally, if an event occurs after hours, we assumed that it took effect during the first trading hour of the next trading day. The same rule applied for news released during weekends or public holidays. The case of global, regional, and local macroeconomic news is explained in the next section.

[39] Other studies such as Dooley and Hutchison (2009); Frank and Hesse (2009); Felices and Wieladek (2012) use similar dates for the GFC.

Table 5.3: Global Financial Crisis timeline and timestamp of events in Eastern Standard Time (EST)

Date	Timestamp	Global financial crisis events
24/07/2007	08:00:00 a.m.	Countrywide Financial Corporation warns of "difficult conditions"
09/08/2007	02:45:00 p.m.	BNP Paribas freezes three funds after being unable to value subprime-mortgage-based assets
13/09/2007	06:23:00 p.m.	Northern Rock receives emergency loan from the Bank of England
16/10/2007	11:37:00 a.m.	Citigroup begins a string of major bank writedowns based on subprime mortgage losses
27/11/2007	11:55:00 p.m.	Citigroup raises $7.5 bn from the Abu Dhabi Investment Authority
11/01/2008	05:02:00 p.m.	Bank of America announces purchase of Countrywide Financial for $4 bn
29/01/2008	01:17:00 p.m.	Rating agencies threaten to downgrade Ambac Financial and MBIA, two major bond insurers
17/02/2008	04:11:00 p.m.	Britain nationalises Northern Rock
13/03/2008	09:46:00 a.m.	Bear Stearns reports a $15 bn (88%) drop in liquid assets
14/03/2008	01:41:00 a.m.	Bear Stearns receives emergency lending from the Fed via J.P. Morgan
16/03/2008	09:30:00 p.m.	J.P. Morgan announces it will purchase Bear Stearns for $2/share
24/03/2008	10:21:00 a.m.	J.P. Morgan's purchase price for Bear Stearns increases to $10/share
06/06/2008	02:21:00 p.m.	S&P downgrades the two largest monoline bond insurers from AAA to AA
16/06/2008	05:06:00 a.m.	Lehman reports a loss of $2.8 bn in the second quarter
11/07/2008	06:08:00 p.m.	After FDIC take-over, IndyMac experiences a run on deposits
10/09/2008	10:05:00 a.m.	Lehman announces $3.9 bn loss in Q3
12/09/2008	10:27:00 a.m.	Moody's and S&P threaten to downgrade Lehman
14/09/2008	11:11:00 p.m.	10 banks create $70 bn liquidity fund
15/09/2008	08:15:00 a.m.	Bank of America purchases Merrill Lynch
15/09/2008	12:41:00 p.m.	Lehman files for bankruptcy
15/09/2008	09:23:00 p.m.	AIG debt downgraded by all three major ratings agencies
16/09/2008	04:58:00 p.m.	RMC money market fund "breaks the buck"
17/09/2008	09:23:00 a.m.	More money market funds come under pressure
25/09/2008	09:07:00 p.m.	WaMu closed by OTS
29/09/2008	08:17:00 a.m.	Systemic risk exception allows open bank assistance to Wachovia
03/10/2008	07:00:00 a.m.	Wells Fargo makes counteroffer for Wachovia
14/10/2008	08:30:00 a.m.	Nine large banks agree to capital injection from the Treasury
23/10/2008	08:16:00 a.m.	Alan Greenspan testifies before the House Committee of Government Oversight and Reform
28/10/2008	11:53:00 p.m.	Consumer confidence hits lowest point on record
30/10/2008	08:30:00 a.m.	Government data shows a 0.3% decline in real US GDP for Q3 2008
23/11/2008	11:50:00 p.m.	Fed, FDIC, and Treasury agree to non-recourse loan to Citigroup if necessary
01/12/2008	12:05:00 p.m.	NBER declares that a recession began in December 2007
17/12/2008	08:30:00 a.m.	November data show a decline in US consumer prices of 1.7%
20/12/2008	03:00:00 a.m.	Eleven of the world's largest banks are downgraded by S&P
10/01/2009	12:01:00 a.m.	US unemployment rises to 7.2%
16/01/2009	08:02:00 a.m.	Citigroup announces plan to split into two units after Q4 loss
24/01/2009	04:38:00 p.m.	Citigroup sells $12 bn of government guaranteed bonds
10/02/2009	11:12:00 a.m.	Markets decline after Geithner's speech due to a lack of specifics
02/03/2009	06:15:00 a.m.	AIG announces $61.7 bn Q4 loss, the largest in US corporate history
02/04/2009	10:52:00 a.m.	Financial Accounting Standards Board (FASB) relaxes mark-to-market accounting rules
09/04/2009	08:14:00 a.m.	Wells Fargo reports record profit in Q1 2009
13/04/2009	04:14:00 p.m.	Goldman Sachs moves to raise $5 bn to pay back TARP funding
05/05/2009	06:42:00 a.m.	Evidence of easing term funding conditions comes as Libor falls below 1%
11/05/2009	02:06:00 p.m.	Following stress test results, banks raise $7.5 bn in new capital
12/05/2009	12:04:00 a.m.	Bank of America sells its stake in China Construction Bank for $7.3 billion
29/05/2009	10:44:00 a.m.	Government statistics show an annualised drop in GDP of 5.7% for Q1 2009
01/06/2009	07:58:00 a.m.	General Motors declares bankruptcy
08/06/2009	05:38:00 a.m.	Ireland's credit rating is cut for the second time in three months
16/07/2009	01:02:00 a.m.	Talks between CIT and government agencies fail to yield a support package
20/07/2009	07:36:00 p.m.	CIT announces $3 bn bond deal and restructuring
07/08/2009	02:00:00 a.m.	Two exchanges agree to end "flash orders" after statement by SEC about potential regulation
25/09/2009	02:05:00 p.m.	Trade volumes for July rose at fastest rate in over five years
01/11/2009	03:39:00 p.m.	CIT Group files for bankruptcy with support of debt holders
13/11/2009	10:00:00 a.m.	Federal Housing Finance Agency capital reserves fall to 0.53%
26/11/2009	12:57:00 a.m.	Dubai World requests six-month debt standstill
09/12/2009	02:56:00 p.m.	Bank of America repays TARP funds
23/11/2009	11:12:00 a.m.	Citibank and Wells Fargo repay TARP funds
27/01/2010	10:44:00 a.m.	Secretary Geithner testifies to congress on AIG deal
10/02/2010	04:31:00 p.m.	PNC Bank repays TARP funds
12/03/2010	02:26:00 a.m.	Examiner's report on the Lehman Bros. Bankruptcy filing released
16/04/2010	10:42:00 a.m.	SEC charges Goldman Sachs with fraud
19/04/2010	08:04:00 a.m.	Citigroup posts $4.4 bn in 2010 Q1 earnings
21/04/2010	11:20:00 a.m.	GM repays remaining TARP funds
06/05/2010	02:20:00 p.m.	The Dow plummets 998.5 points, its largest intraday point drop ever
25/11/2010	10:00:00 a.m.	New home sales hit lowest levels on record
27/11/2010	10:00:00 a.m.	Stocks rally following Chairman Bernanke's speech
29/09/2010	05:56:00 p.m.	JPM and GMAC freeze foreclosures proceedings

Date	Timestamp	Global financial crisis events
30/09/2010	12:00:00 a.m.	AIG announces government assistance exit plan
01/10/2010	04:57:00 p.m.	Bank of America freezes foreclosures proceedings
17/11/2010	04:54:00 p.m.	GM has a $20.1 bn IPO, the largest in US history
01/12/2010	05:26:00 p.m.	AIG issues its first bond since its near collapse
17/12/2010	01:50:00 p.m.	B of A is sued for routinely misleading consumers about home loan modifications
27/12/2010	09:12:00 a.m.	AIG obtains new credit lines from commercial banks to replace its FRBNY bailout aid
28/12/2010	08:30:00 a.m.	Initial claims for unemployment fell to their lowest level in 2 years

Source: Authors' compilation of market events outlined in the financial turmoil timeline of the Federal Reserve of New York (FEDNewYork, 2011) with the exception of the first event which is taken from the crisis timeline of the Federal Reserve of St. Louis (FEDSt.Louis, 2009).

5.3 Method

Based on findings of other studies,[40] we use a common set of macroeconomic news items that have been found to have a significant impact on government bonds. We treat events derived from the GFC and macroeconomic announcements from the US and Europe as global events, macroeconomic announcements from Brazil and Mexico as regional,[41] and macroeconomic announcements from Colombia as local. In line with the event study method, we want to see if a surprise (in the case of macroeconomic news) has a significant effect on asset prices in the form of positive or negative abnormal returns. In the case of the GFC-sourced events (see Table 5.3) we assume that all events are surprises. For this study, all the macroeconomic news announcements were collected from Bloomberg and we calculate the standardised surprise component in the same way as Balduzzi et al. (2001) as in equation (5.1):

$$S_{i,t} = \frac{A_{i,t} - F_{i,t}}{\sigma_{i,t}} \tag{5.1}$$

In Equation 5.1, $A_{i,t}$ = the actual value of the announcement *i* at time *t*, $F_{i,t}$ = the average forecasted value of the announcement *i* at time *t*, which in the case of Bloomberg data is the average value of all the forecasted values of the analysts that contribute to the system, or in other words this is a proxy of market consensus, $\sigma_{i,t}$ = the standard deviation obtained from the time series of the difference between the actual value and the forecast $(A_{i,t} - F_{i,t})$ of each a macroeconomic announcement *i* at time *t* and S_i = the standardised surprise component of an announcement *i* at time *t*.

[40] The classical examples in the literature can be found in Fleming and Remolona (1999), Balduzzi et al. (2001), and Andersen et al. (2007). More recent examples on the effects of macroeconomic news on bonds can be found in Aizenman et al. (2012) and Beber and Brandt (2010).

[41] This is based on anecdotal evidence provided by analysts from brokerage firms in Colombia.

In Table 5.4 we summarise the different macroeconomic announcements and their level of surprise in number of standard deviations from the mean, where 0 standard deviations can be interpreted as in line with the consensus and ±3 standard deviations represents an extreme surprise.

Table 5.4: Number of global, regional, and local macroeconomic announcements from 2007 to 2013 and level of surprise

	Country	Ticker	News name	Level of surprise in standard deviations							Total
				-3	-2	-1	0	1	2	3	
GLOBAL NEWS	GFC	GFC	Global events derived from the GFC (see Table 3)							74	74
	US	CPI CHNG	Consumer Price Index (MoM)		3	5	67	6	2	1	84
		CPUPXCHG	CPI Ex Food & Energy (MoM)		1	20	37	23	3		84
		FDTR Inde	FOMC Rate Decision			3	49	1	1	4	58
		GDP CQOQ	GDP QoQ (Annualised)	1	1	8	63	8	3		84
		GDP PIQQ	GDP Price Index		2	6	65	6	5		84
		INJCJC In	Initial Jobless Claims	3	12	30	267	40	4	2	358
		NFP TCH I	Change in Nonfarm Payrolls		1	13	54	15	1		84
		PPI CHNG	Producer Price Index (MoM)	1		11	66	3	3		84
		RSTAO I	Adjusted Retail & Food Services Sales SA Total Monthly % Change	1		12	60	8	3		84
		USCABAL I	Current Account Balance			5	20	2	1		28
		USTBTOT I	Trade Balance		3	10	55	14	2		84
		USURTOT I	Unemployment Rate		2	11	53	15	3		84
		CONSSENT	U. of Michigan Confidence		4	11	125	17	7		164
		WINCHNG	Merchant Wholesalers Inventories Total Monthly % Change		1	13	58	10	2		84
	Eurozone	ECCPEST I	Eurozone CPI Estimate (YoY)		4	7	61	8	4		84
		EUPPEU	Eurostat PPI Eurozone Industry Ex Construction MoM		1	13	59	8	2	1	84
		EUPPEUY	Eurostat PPI Eurozone Industry Ex Construction YoY		1	5	69	6	1	2	84
		EURR002W	ECB Announces Interest Rates	1	1	2	73	2	1	3	83
		EUSATOTN	ECB Eurozone Current Account SA			2	7	2		73	84
		PIEZCA	Markit Eurozone Composite PMI SA	1			28	2		126	157
		PITEZ I	PMI Manufacturing			14	56	11	2	1	84
		RSSAEU	Eurostat Retail Sales Volume Eurozone MoM SA		1	11	59	10	3		84
		URTEU I	Eurostat Unemployment Eurozone SA	1	1	6	61	12	2	1	84
		XTTBEZ In	Eurozone Trade Balance			13	51	9	3	9	85
		ECCPEUY	Eurostat European Union HICP All Items YoY NSA		7		67		10		84
F F	Mexico	IEFANF	General Manufacturing Index Mexico			7	35	5		22	69
		IEFNAN	General Non Manufacturing Index Mexico		1	6	32	5	1	23	68
		XBWCORE	Mexico CPI Core inflation Percent Change Biweekly	1	2	10	76	12	2	3	106
		XBWO I	Mexico CPI Change with respect with previous observation Biweekly		2	14	76	9	5	1	107
		XCACUAC	Mexico Nominal Current Account Balance		2	14	76	9	5	1	107
		XCPCHNG	Mexico CPI MoM		1	15	53	11	4		84
		XCPYOY I	Mexico CPI YoY		2	14	76	9	5	1	107
		XGCTOT I	Mexico GDP Total YoY NSA 2008=100			7	18	3			28
		XIPTYOY	MX Industrial Production Total Yearly % Change		3	9	58	10	2		82
		XONBR In	Bank of Mexico Official Overnight Rate		3	1	60	2	1	2	69
		XTBBAL I	Mexico Trade Balance Monthly Total USD Million		2	18	73	9	5	51	158
		XUERATE	Mexico Unemployment Rate for Workers 14 and Older ENOE NSA		2	13	56	11	2		84
		XVPTOTL	Mexico Vehicle Production Total Production		1	1	10	2		32	46
		XWRTRYO	Mexico Wholesale/Retail Sale YOY Total Return NSA		2	10	61	7	4		84
	Brazil	BZCACURR	Current Account - Monthly		1	8	62	8	4	1	84

				Level of surprise in standard deviations							
	Country	**Ticker**	**News name**	-3	-2	-1	0	1	2	3	**Total**
		BZDPNDT	Net Debt % GDP	1	1		77	4		1	84
		BZGDQOQ	GDP (IBGE) QoQ		1	3	21	2	1		28
		BZIPYOY	Industrial Production YoY			14	60	8		2	84
		BZPIIPC	Brazil CPI		3	7	60	10	3	1	84
		BZRTRET	Brazil Retail Sales Volume		2	8	44	4	3	3	64
		BZRTRYOY	Retail Sales (YoY)		3	7	63	9	1	1	84
		BZSTSETA	SELIC Target - Central Bank		1	4	42	7	1	1	56
		BZTBEXP	Brazil Trade Balance FOB Exports			1	82			1	84
		BZTBIP	Brazil Trade Balance FOB Imports NSA	1	1	9	62	8	2	1	84
		BZUETOTN	Unemployment Rate		2	16	58	6	1	1	84
		IBREGPD	FGV Brazil General Prices IGP-DI MoM		2	6	65	6	5		84
LOCAL NEWS	Colombia	COCIPIBY	GDP (YoY)		1	1	19	2	1	1	25
		COCPIO	Colombian Inflation Old index		2	10	59	12	1		84
		COCPIYOY	Consumer Price Index (YoY)		2	10	58	13	1		84
		CODRYOY I	GDP (YoY)				2	1		2	5
		COIPEYOY	Industrial Production (YoY)		1	5	69	5	3	1	84
		CONCCONF	Consumer Confidence		2	12	54	15	1		84
		CORRRIN	Colombia Minimum Repo Rate to Be Offered at the Daily Auction			7	68	4	2	3	84
		COSAYOY I	Retail Sales (YoY)	1	1	10	62	9	1		84
		COTRBAL	Colombia Trade Balance FOB	1	2	3	46	11		22	85
		COUNTOTR	Urban Unemployment Rate	1	1	7	55	14		6	84
		Total general		15	98	518	3578	500	135	481	5325

Source: Authors' compilation of data extracted from Bloomberg, the level of the surprise is estimated using equation ($S_{i,t} = \frac{A_{i,t} - F_{i,t}}{\sigma_{i,t}}$) where $A_{i,t}$ = the actual value of the announcement *i* at time *t*, $F_{i,t}$ = the average forecasted value of the announcement *i* at time *t*, which in the case of Bloomberg data is the average value of all the forecasted values of the analysts that contribute to the system, or in other words this a proxy of market consensus, $\sigma_{i,t}$ = the standard deviation for the series sample of a specific macroeconomic announcement *i* at time *t* and S_i = the standardised surprise component of an announcement *i* at time *t*. The level of surprise is in number of standard deviations from the mean where 0 standard deviations can be interpreted as in line with the consensus and +/-3 standard deviations represents an extreme surprise. As in Balduzzi et al.(2001) and for the remainder of the paper, we categorise as a surprise all the observations with a level of +/-1 standard deviations from the consensus (mean).

The next step is to model the correlation between the Colombian government bonds and the US market factor using the factor model in equation (5.2):

$$R_{i,t} = \alpha_i + \beta_i R_{m,t} + e_{i,t} \tag{5.2}$$

where $R_{i,t}$ =the compounded hourly return of the Colombian bonds *i* at time *t*, $R_{m,t}$ = the compounded hourly return of the S&P 500 ETF, and $e_{i,t}$ = the idiosyncratic error term. $e_{i,t}$ is equivalent to the abnormal return since equation (5.2) can be rearranged into equation (5.3):

$$e_{i,t} = R_{i,t} - (\alpha_i + \beta_i R_{m,t}) \tag{5.3}$$

Therefore, we can determine if an abnormal return *i* at time *t* is significant due to additional information in an unanticipated surprise or event that is not captured by the market model:

$$t - stat = \frac{e_{i,t}}{\sigma_{est}} \tag{5.4}$$

where σ_{est} = the standard error of the regression obtained from estimating equation (5.2) using all hourly observations that do not contain an event and then predicting the returns of the excluded observations that contain an event. Therefore, if we reject the null that $e_{i,t}$ of an event is not statistically different to zero we can infer the significance of the event *i* at time *t*. We can identify GFC, global, regional, and local significant events based on the level of the surprise associated with an event as calculated by equation (5.1). If there is decoupling from GFC events, then these events should not be significant and with a greater impact than other surprises originating from normal macroeconomic announcements.

Finally, we use the specification suggested by Balduzzi et al. (2001) to measure the effect of a particular macroeconomic surprise on bond returns. For this purpose we run a regression of the surprises on the bond returns using the specification in equation (5.5):

$$e_{i,t} = \alpha_i + \beta_i S_{i,t} + \varepsilon_{i,t} \tag{5.5}$$

where $e_{i,t}$ = the abnormal return, $S_{i,t}$ = all the standardised level of surprises from Table 4.4 that are equal to or greater than ±1 standard deviations from the event sample mean where 0 standard deviations can be interpreted as in line with the consensus and ±3 standard deviations represents an extreme surprise, and β_i measures the sensitivity of returns to that particular event. In the next section we summarise the results. Each surprise is matched with the abnormal return of any issue that traded at the same time of the announcement, or in the case of an after-hours surprise, we match it with the abnormal return of the issue that traded in the first hour of the next trading day.

5.4 Results

In Table 5.5 we summarise the results obtained from applying the market model in equation (4.2).

Table 5.5: Summary of significant results obtained from the market model

Issue	Coefficient	Error	T-Stat	p-value	N	R^2
TFIT07220808	0.005*	(0.0026)	1.865	0.0623	1147	0.303%
TFIT03140509	0.003*	(0.0016)	1.878	0.0606	1690	0.208%
TFIT05100709	0.011**	(0.0050)	2.136	0.0329	1032	0.441%
TFIT05241110	0.006**	(0.0026)	2.299	0.0215	3049	0.173%
TFIT04180511	0.008***	(0.0024)	3.230	0.0125	2596	0.401%
TFIT10281015	0.105***	(0.0121)	8.718	0.0000	3488	2.134%
TFIT11241018	0.073***	(0.0213)	3.447	0.0050	1949	0.607%
TFIT15240720	0.066***	(0.0083)	7.977	0.0000	5227	1.203%
TFIT10040522	0.089***	(0.0336)	2.638	0.0084	1377	0.503%

The results in this table are obtained from running the market model regression $R_{i,t} = \alpha_i + \beta_i R_{m,t} + e_{i,t}$ where $R_{i,t}$ = the compounded hourly return of the Colombian bonds *i* at time *t*, $R_{m,t}$ = the compounded hourly return of the S&P 500 ETF, and $e_{i,t}$ = the idiosyncratic error term. In these regressions we exclude the hours without relevant news or events from the data as well as missing contemporaneous observations from daylight savings time. In this table we report those bonds with statistically significant beta coefficients obtained by running the regression obtained from equation (5.2).

For nine of the 16 Colombian bonds that we analysed, the US market factor has some explanatory power, although with very low R^2. Interestingly, the issues for which the market factor is significant are those with higher coupon rates. This is in contrast with findings in previous studies (see: Dick-Nielsen et al., 2012; Friewald et al., 2012; Pelizzon et al., 2013), where, guided by the common belief that issues with lower coupon rates are the ones that should reflect lower credit risk, lower coupon issues are found to have higher liquidity and higher sensitivity to market influences. Therefore, we hypothesise that Colombian local-currency bond market traders prefer bonds with higher coupons and higher sensitivity to price changes for speculative purposes.

In Table 5.6 we summarise the effects of the surprises on issues with significant market factor sensitivity from Table 5.5.

Table 5.6: Effect of macroeconomic surprises on Colombian bonds

		TFIT07220808			TFIT03140509			TFIT05100709			TFIT05241110		
		Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2
GLOBAL NEWS	Consumer Price Index (MoM)												
	CPI Ex Food & Energy (MoM)										0.043	(0.195)	0.60%
	FOMC Rate Decision												
	GDP QoQ (Annualised)										0.003	(0.016)	0.26%
	GDP Price Index										-0.052***	(0.011)	70.60%
	Initial Jobless Claims				0.000	(0.000)	2.96%				0.000	(0.000)	2.97%
	Change in Nonfarm Payrolls	0.000	(0.000)	24.39%							0.000**	(0.000)	52.64%
	Producer Price Index (MoM)										-0.099*	(0.043)	46.96%
	Adjusted Retail & Food Services Sales SA Total Monthly % Change										-0.014*	(0.008)	31.66%
	Current Account Balance												
	Trade Balance										0.000	(0.000)	5.39%
	Unemployment Rate										0.031	(0.261)	0.27%
	U. of Michigan Confidence	0.000	(0.000)	49.40%	0.000	(0.000)	5.93%	0.000	(0.000)	46.94%	0.000	(0.000)	0.17%
	Merchant Wholesalers Inventories Total Monthly % Change				0.036**	(0.012)	53.98%				0.077**	(0.035)	25.80%
	Eurozone CPI Estimate (YoY)				-0.014	(0.050)	1.16%				0.157	(0.160)	19.52%
	Eurostat PPI Eurozone Industry Ex Construction MoM												
	Eurostat PPI Eurozone Industry Ex Construction YoY												
	ECB Announces Interest Rates												
	ECB Eurozone Current Account SA							0.000***	(0.000)	100.00%			
	Markit Eurozone Composite PMI SA												
	PMI Manufacturing				0.000	(0.000)	41.79%				0.000	(0.000)	0.77%
	Eurostat Retail Sales Volume Eurozone MoM SA				-0.034**	(0.011)	68.44%				-0.140**	(0.045)	70.35%
	Eurostat Unemployment Eurozone SA	-0.020***	(0.000)	100.00%									
	Eurozone Trade Balance	0.000	(0.000)	9.62%	0.000	(0.000)	14.08%				0.000	(0.000)	16.60%
	Eurostat European Union HICP All Items YoY NSA				0.000	(0.027)	0.00%				-0.187***	(0.014)	97.30%
R E	General Manufacturing Index Mexico												
	General Non Manufacturing Index Mexico												
	Mexico CPI Core inflation Percent Change Biweekly	0.078***	(0.008)	95.68%	-0.275**	(0.100)	36.79%	0.538	(1.352)	3.81%	-0.115	(0.129)	5.07%
	Mexico CPI Change with respect with previous observation Biweekly												
	Mexico Nominal Current Account Balance				0.282	(0.156)	26.78%						
	Mexico CPI MoM										-0.551	(0.378)	19.10%
	Mexico CPI YoY												
	Mexico GDP Total YoY NSA 2008=100												
	MX Industrial Production Total Yearly % Change										0.120	(0.043)	79.25%
	Bank of Mexico Official Overnight Rate				-0.012	(0.016)	14.91%				-0.010	(0.015)	13.10%
	Mexico Trade Balance Monthly Total USD Million				0.000**	(0.000)	81.24%				0.000	(0.000)	63.06%
	Mexico Unemployment Rate for Workers 14 and Older ENOE NSA										-0.027	(0.011)	75.53%
	Mexico Vehicle Production Total Production												
	Mexico Wholesale/Retail Sale YOY Total Return NSA										0.011	(0.008)	20.87%
	Current Account - Monthly-Brazil												
	Net Debt % GDP												

		TFIT07220808			TFIT03140509			TFIT05100709			TFIT05241110		
		Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2
	GDP (IBGE) QoQ												
	Industrial Production YoY										0.045	(0.025)	35.82%
	Brazil CPI										0.446	(0.590)	16.02%
	Brazil Retail Sales Volume										-0.075	(0.039)	38.23%
	Retail Sales (YoY)										0.007*	(0.003)	44.88%
	SELIC Target - Central Bank										0.218	(0.117)	41.01%
	Brazil Trade Balance FOB Exports												
	Brazil Trade Balance FOB Imports NSA										0.000	(0.000)	21.45%
	Unemployment Rate-Brazil										0.031	(0.261)	0.27%
	FGV Brazil General Prices IGP-DI MoM										0.025***	(0.007)	70.01%
LOCAL NEWS	GDP (YoY)												
	Colombian Inflation Old index												
	Consumer Price Index (YoY)												
	Industrial Production (YoY)				-0.018**	(0.002)	96.86%				-0.048	(0.021)	72.02%
	Consumer Confidence				0.000	(0.000)	16.36%				0.000	(0.000)	11.87%
	Colombia Minimum Repo Rate to Be Offered at the Daily Auction				0.480***	(0.030)	98.50%				1.280***	(0.108)	95.21%
	Retail Sales (YoY)										0.007*	(0.003)	44.88%
	Colombia Trade Balance FOB												
	Urban Unemployment Rate-Colombia				0.000	(0.004)	0.06%				0.001	(0.007)	0.39%

The results in this table are obtained from running the regression $e_{i,t} = \alpha_i + \beta_i S_{i,t} + \varepsilon_{i,t}$ where $e_{i,t}$ = the abnormal returns from equation (5.3) , $S_{i,t}$ = all the standardised surprises that are equal or greater to one standard deviation from the event sample mean, and finally β_i measures the sensitivity of returns to that particular event. Each surprise is matched with any issue that traded at the same time of the announcement, or in the case of an after-hours surprise we match it with the abnormal return of the issue that traded in the first hour of the next trading day. Therefore, that is why expired and less liquid issues have fewer matches than on-the run or more liquid issues.

Table 5.6: Effect of macroeconomic surprises on Colombian bonds (cont.)

		TFIT04180511			TFIT10281015			TFIT11241018			TFIT15240720		
		Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2
GLOBAL NEWS	Consumer Price Index (MoM)										-0.397***	(0.042)	93.75%
	CPI Ex Food & Energy (MoM)	-0.645**	(0.280)	36.99%	-0.081	(0.217)	0.64%	-1.969**	(0.711)	46.02%	0.356	(1.127)	0.52%
	FOMC Rate Decision												
	GDP QoQ (Annualised)	0.124**	(0.043)	50.89%	0.128	(0.079)	22.41%				0.187	(0.191)	7.41%
	GDP Price Index							0.054**	(0.019)	61.29%	-0.067	(0.041)	28.27%
	Initial Jobless Claims	0.000	(0.000)	3.27%	0.000	(0.000)	2.18%	0.000*	(0.000)	30.01%	0.000	(0.000)	3.72%
	Change in Nonfarm Payrolls	0.000***	(0.000)	47.34%	0.000	(0.000)	3.05%	0.000	(0.000)	33.91%	0.000	(0.000)	18.57%
	Producer Price Index (MoM)	-0.036	(0.051)	5.83%				0.201***	(0.013)	98.74%	0.143	(0.197)	3.59%
	Adjusted Retail & Food Services Sales SA Total Monthly % Change	0.026	(0.033)	7.25%	0.421***	(0.038)	98.43%	0.081	(0.052)	54.56%	0.013	(0.178)	0.05%
	Current Account Balance										0.000*	(0.000)	89.77%
	Trade Balance	0.000	(0.000)	4.05%	0.000	(0.000)	0.44%				0.000	(0.000)	10.00%
	Unemployment Rate	-0.235	(0.165)	28.89%	-0.367	(0.227)	22.61%				-0.147	(0.385)	0.85%
	U. of Michigan Confidence	0.000	(0.000)	0.00%	0.000	(0.000)	1.36%	0.000	(0.001)	0.99%	0.000*	(0.000)	11.24%
	Merchant Wholesalers Inventories Total Monthly % Change	0.251*	(0.120)	25.13%	0.019	(0.062)	0.89%	0.043	(0.034)	44.80%	-0.015	(0.065)	0.56%
	Eurozone CPI Estimate (YoY)	0.146**	(0.031)	91.96%							0.360**	(0.148)	22.84%
	Eurostat PPI Eurozone Industry Ex Construction MoM	-0.219***	(0.023)	92.14%	-0.005	(0.071)	0.06%	-0.027	(0.043)	8.63%	0.458	(0.681)	6.05%
	Eurostat PPI Eurozone Industry Ex Construction YoY												
	ECB Announces Interest Rates												
	ECB Eurozone Current Account SA												
	Markit Eurozone Composite PMI SA												
	PMI Manufacturing	0.000	(0.000)	20.51%	0.000	(0.000)	5.30%	0.000	(0.001)	3.02%	0.000	(0.000)	5.45%
	Eurostat Retail Sales Volume Eurozone MoM SA	-0.002	(0.120)	0.00%	-0.006	(0.051)	0.19%	0.028	(0.109)	1.32%	0.083	(0.094)	16.41%
	Eurostat Unemployment Eurozone SA	0.088	(0.154)	3.97%	-0.398	(0.739)	6.78%				-1.515*	(0.793)	21.94%
	Eurozone Trade Balance	0.000	(0.000)	50.30%	0.000	(0.000)	37.29%	0.000	(0.000)	0.31%	0.000*	(0.000)	16.62%
	Eurostat European Union HICP All Items YoY NSA				0.155	(0.295)	8.40%				-0.157	(0.497)	1.41%
REGIONAL NEWS	General Manufacturing Index Mexico				0.000***	(0.000)	68.52%	0.000	(0.000)	1.45%	0.001	(0.000)	20.59%
	General Non Manufacturing Index Mexico							0.000*	(0.000)	32.68%	-0.001**	(0.000)	76.38%
	Mexico CPI Core inflation Percent Change Biweekly				-0.245	(0.148)	18.70%	0.400**	(0.085)	91.75%	0.378	(0.418)	4.33%
	Mexico CPI Change with respect with previous observation Biweekly	-0.141***	(0.004)	98.90%	0.204***	(0.053)	47.82%				-0.262	(0.416)	3.80%
	Mexico Nominal Current Account Balance				-0.510	(0.297)	21.15%	0.275***	(0.005)	99.57%	-0.928**	(0.356)	36.17%
	Mexico CPI MoM	0.003	(0.309)	0.00%				0.115	(1.229)	0.29%	-1.371***	(0.475)	25.80%
	Mexico CPI YoY	-0.164*	(0.081)	33.88%	-0.641***	(0.210)	39.91%				0.097	(0.172)	1.73%
	Mexico GDP Total YoY NSA 2008=100												
	MX Industrial Production Total Yearly % Change	-0.007	(0.008)	15.28%	-0.011	(0.016)	13.22%				-0.037	(0.048)	6.07%
	Bank of Mexico Official Overnight Rate										-0.126	(0.439)	2.04%
	Mexico Trade Balance Monthly Total USD Million	0.000*	(0.000)	76.38%	0.000	(0.000)	5.41%	0.000**	(0.000)	44.66%	0.000	(0.000)	0.37%
	Mexico Unemployment Rate for Workers 14 and Older ENOE NSA	-0.044	(0.033)	37.75%	-0.026	(0.071)	2.62%	-0.192	(0.111)	37.46%	-0.186	(0.128)	18.86%
	Mexico Vehicle Production Total Production												
	Mexico Wholesale/Retail Sale YOY Total Return NSA	0.025	(0.016)	43.40%	0.003	(0.014)	0.65%	0.001	(0.028)	0.14%	0.042	(0.030)	14.99%
	Current Account - Monthly	0.000***	(0.000)	54.21%	0.000	(0.000)	36.38%	0.000	(0.000)	1.27%	0.000	(0.000)	1.49%
	Net Debt % GDP												

		TFIT04180511			TFIT10281015			TFIT11241018			TFIT15240720		
		Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2	Surprise Coefficient	Error	R^2
	GDP (IBGE) QoQ										0.118	(0.076)	25.58%
	Industrial Production YoY	0.076***	(0.022)	56.83%							-0.117	(0.073)	29.93%
	Brazil CPI				0.087	(0.123)	2.31%				-1.025*	(0.550)	27.87%
	Brazil Retail Sales Volume				-0.069***	(0.005)	96.56%	-0.019	(0.030)	8.95%	-0.079	(0.046)	29.49%
	Retail Sales (YoY)	0.019	(0.015)	13.64%	0.018*	(0.008)	46.72%				-0.036	(0.034)	10.81%
	SELIC Target - Central Bank	0.045	(0.160)	0.66%							2.632*	(1.301)	29.04%
	Brazil Trade Balance FOB Exports												
	Brazil Trade Balance FOB Imports NSA	0.000	(0.000)	1.69%	0.000	(0.000)	10.70%	0.000	(0.000)	0.38%	0.000***	(0.000)	42.05%
	Unemployment Rate	-0.235	(0.165)	28.89%	-0.367	(0.227)	22.61%				-0.147	(0.385)	0.85%
	FGV Brazil General Prices IGP-DI MoM	0.087***	(0.005)	97.89%				-0.106***	(0.009)	95.75%	-0.094	(0.069)	20.90%
LOCAL NEWS	GDP (YoY)												
	Colombian Inflation Old index										1.022**	(0.407)	41.19%
	Consumer Price Index (YoY)				-2.469***	(0.061)	99.34%				-5.252**	(2.012)	25.42%
	Industrial Production (YoY)	-0.013*	(0.006)	47.83%	0.054	(0.026)	68.87%				-0.175	(0.096)	52.30%
	Consumer Confidence	0.000	(0.000)	21.17%	0.000	(0.000)	0.09%	0.000	(0.000)	7.18%	0.000**	(0.000)	15.97%
	Colombia Minimum Repo Rate to Be Offered at the Daily Auction	0.870***	(0.193)	80.28%	2.331***	(0.335)	87.37%	2.343***	(0.309)	91.99%	2.029***	(0.541)	70.09%
	Retail Sales (YoY)	0.019	(0.015)	13.64%	0.018*	(0.008)	46.72%				-0.036	(0.034)	10.81%
	Colombia Trade Balance FOB				0.000	(0.000)	3.18%				0.000**	(0.000)	55.62%
	Urban Unemployment Rate	0.004*	(0.002)	34.39%	0.044*	(0.022)	39.07%	-0.040	(0.027)	42.77%	0.010	(0.023)	1.92%

The results in this table are obtained from running the regression $e_{i,t} = \alpha_i + \beta_i S_{i,t} + \varepsilon_{i,t}$ where $e_{i,t}$ = the abnormal returns from equation (5.3), $S_{i,t}$ = all the standardised surprises that are equal to or greater than one standard deviation from the event sample mean, and finally β_i measures the sensitivity of returns to that particular event. Each surprise is matched with any issue that traded at the same time of the announcement, or in the case of an after-hours surprise we match it with the abnormal return of the issue that traded in the first hour of the next trading day. Therefore, that is why expired and less liquid issues have fewer matches than on-the run or more liquid issues.

Table 5.6: Effect of macroeconomic surprises on Colombian bonds (cont.)

		TFIT10040522		
		Surprise Coefficient	Error	R^2
GLOBAL NEWS	Consumer Price Index (MoM)			
	CPI Ex Food & Energy (MoM)	-1.038*	(0.511)	27.25%
	FOMC Rate Decision			
	GDP QoQ (Annualised)	0.622	(0.753)	25.47%
	GDP Price Index			
	Initial Jobless Claims	0.000	(0.000)	11.91%
	Change in Nonfarm Payrolls			
	Producer Price Index (MoM)			
	Adjusted Retail & Food Services Sales SA Total Monthly % Change			
	Current Account Balance			
	Trade Balance	0.000	(0.001)	5.17%
	Unemployment Rate			
	U. of Michigan Confidence	0.000*	(0.000)	51.47%
	Merchant Wholesalers Inventories Total Monthly % Change	0.015	(0.093)	1.28%
	Eurozone CPI Estimate (YoY)	-1.314	(0.970)	47.85%
	Eurostat PPI Eurozone Industry Ex Construction MoM			
	Eurostat PPI Eurozone Industry Ex Construction YoY			
	ECB Announces Interest Rates			
	ECB Eurozone Current Account SA			
	Markit Eurozone Composite PMI SA			
	PMI Manufacturing			
	Eurostat Retail Sales Volume Eurozone MoM SA			
	Eurostat Unemployment Eurozone SA			
	Eurozone Trade Balance			
	Eurostat European Union HICP All Items YoY NSA			
REGIONAL NEWS	General Manufacturing Index Mexico			
	General Non Manufacturing Index Mexico			
	Mexico CPI Core inflation Percent Change Biweekly	-2.540*	(1.179)	39.86%
	Mexico CPI Change with respect with previous observation Biweekly			
	Mexico Nominal Current Account Balance			
	Mexico CPI MoM			
	Mexico CPI YoY			
	Mexico GDP Total YoY NSA 2008=100			
	MX Industrial Production Total Yearly % Change			
	Bank of Mexico Official Overnight Rate			
	Mexico Trade Balance Monthly Total USD Million	0.000	(0.000)	0.03%
	Mexico Unemployment Rate for Workers 14 and Older ENOE NSA	0.183	(0.513)	5.97%
	Mexico Vehicle Production Total Production			
	Mexico Wholesale/Retail Sale YOY Total Return NSA	0.031	(0.046)	10.23%
	Current Account - Monthly			
	Net Debt % GDP			

		Surprise Coefficient	Error	R^2
	GDP (IBGE) QoQ			
	Industrial Production YoY			
	Brazil CPI	-5.942	(20.172)	2.81%
	Brazil Retail Sales Volume			
	Retail Sales (YoY)	0.052	(0.061)	19.60%
	SELIC Target - Central Bank			
	Brazil Trade Balance FOB Exports			
	Brazil Trade Balance FOB Imports NSA	0.000	(0.000)	12.89%
	Unemployment Rate			
	FGV Brazil General Prices IGP-DI MoM			
LOCAL NEWS	GDP (YoY)			
	Colombian Inflation Old index			
	Consumer Price Index (YoY)			
	Industrial Production (YoY)			
	Consumer Confidence	0.000	(0.000)	0.01%
	Colombia Minimum Repo Rate to Be Offered at the Daily Auction			
	Retail Sales (YoY)	0.052	(0.061)	19.60%
	Colombia Trade Balance FOB	0.000	(0.000)	1.52%
	Urban Unemployment Rate			

The results in this table are obtained from running the regression $e_{i,t} = \alpha_i + \beta_i S_{i,t} + \varepsilon_{i,t}$ where $e_{i,t}$ = the abnormal returns from equation (5.3), $S_{i,t}$ = all the standardised surprises that are equal to or greater than one standard deviation from the event sample mean, and finally β_i measures the sensitivity of returns to that particular event. Each surprise is matched with any issue that traded at the same time of the announcement, or in the case of an after-hours surprise we match it with the abnormal return of the issue that traded in the first hour of the next trading day. Therefore, that is why expired and less liquid issues have fewer matches than on-the run or more liquid issues.

Table 5.6 shows that twelve (12) types of global macroeconomic news are significant, with coefficients greater than zero. These are: CPI-Consumer Price Index (US), CPI Ex Food and Energy (US), GDP QoQ Annualised (US), GDP Price Index, Producer Price Index-PPI, Adjusted Retail Food and Services (US), Merchant Wholesales Inventories (US), CPI Eurozone, PPI Eurozone, Euro Stat Retail Sales, Unemployment Eurozone, and the Eurozone Harmonised Indices of Consumer prices. Across medium- and long-term maturities the most significant surprises are those related to CPI Ex Food and Energy (US) and different measures of the Eurozone CPI. These effects can be seen in at least one of the most liquid on-the run issues (TFIT52240720) and one expired (but previously liquid) issue (TFIT241110). The highest impact in a particular issue (TFIT22240720) was due to a surprise in Unemployment Eurozone. Other significant surprise coefficients but with a lesser impact were those related to trade such as PPI (US), retail sales (US), retail sales (Eurozone), and Inventories (US).

In the case of regional news, eleven (11) types of news are significant with coefficients greater than zero, although most of them are different measures of inflation. For example, in the case of Mexico: Mexico CPI Core Inflation Percent Change Biweekly, Mexico CPI Change with respect to previous observation Biweekly, Mexico CPI month on month, and Mexico CPI year on year. The only news event for Mexico that was not inflation-related was the Mexico Nominal Current Account Balance. In the case of Brazil: Industrial Production, Brazil Retail Sales Volume, Retail (year on year), Brazil CPI, FGV Brazil General Prices, and SELIC (Interest Rates) Target-Central Bank. The highest impact in a particular issue (TFIT22240720) was due to a surprise in the Brazilian target interest rate (SELIC

Target). It is interesting that in the case of CPI surprises in regional news, the impact of the surprise on average is higher than for global inflation news.

In the case of surprises from local news we can observe seven (7) significant events: Colombia Inflation Old Index, Consumer Price Index (year on year), Industrial Production (year on year), Colombian Minimun Repo Rate to be Offered at Auction (interest rate), and Urban Unemployment Rate. In the case of local news the highest impact across issues is from the repo rate and the highest impact for a particular issue (TFIT22240720) was due to a surprise in the Colombian CPI with a higher impact than any kind of global and regional news.

In Table 5.7 we summarise the average impact that one-standard-deviation significant surprises have on bond prices. Local news events have the highest impact on Colombian bond prices.

Table 5.7: Average effect of the surprises in global, regional, and local news in Colombian bond prices

Global News	Effect
Consumer Price Index (MoM)	-0.057%
CPI Ex Food & Energy (MoM)	-0.102%
GDP QoQ (Annualised)	0.055%
GDP Price Index	0.023%
Producer Price Index (MoM)	0.024%
Adjusted Retail & Food Services Sales SA Total Monthly % Change	0.104%
Merchant Wholesalers Inventories Total Monthly % Change	0.070%
Eurozone CPI Estimate (YoY)	0.038%
Eurostat PPI Eurozone Industry Ex Construction MoM	-0.040%
Eurostat Retail Sales Volume Eurozone MoM SA	-0.040%
Eurostat Unemployment Eurozone SA	-0.080%
Eurostat European Union HICP All Items YoY NSA	-0.008%
Regional News	
Mexico CPI Core inflation Percent Change Biweekly	-0.049%
Mexico CPI Change with respect with previous observation Biweekly	0.004%
Mexico Nominal Current Account Balance	-0.045%
Mexico CPI MoM	-0.080%
Mexico CPI YoY	-0.056%
Industrial Production YoY	0.105%
Brazil CPI	-0.057%
Brazil Retail Sales Volume	-0.049%
Retail Sales (YoY)	0.015%
SELIC Target - Central Bank	0.311%
FGV Brazil General Prices IGP-DI MoM	0.001%
Local News	
Colombian Inflation Old index	0.198%
Consumer Price Index (YoY)	-0.760%
Industrial Production (YoY)	-0.050%
Colombia Minimum Repo Rate to Be Offered at the Daily Auction	0.217%
Retail Sales (YoY)	0.024%
Urban Unemployment Rate	0.017%

The results in this table are obtained by multiplying the significant coefficients in Table 5.6 by the standard deviation of each news series. The final result in the case of news that affect multiple issues is the average of the results obtained for each issue.

In contrast to normal macroeconomic news, the events sourced in the financial crisis cannot be standardised in terms of surprises, since they cannot be defined in terms of level of surprise from a given consensus and are just ad-hoc news that rattled the market. Our approach was to first test for the significance of each abnormal return using the procedure described in equations (5.2) to (5.4). The second step was to match abnormal returns for each bond issue with all the GFC events outlined in Table (5.3) and select those that were significant for a particular date. In Table 5.8 we show the significant abnormal returns for a GFC event as well as the matched US market proxy return for that date. The results highlighted in bold are negative events; all others are positive events.

Table 5.8: Significant effects of GFC events on Colombian bonds abnormal returns

		TFIT07220808		
Number of events	**Date**	**Hour**	**AR**	**IVV Return**
5	24/03/2008	10:21:00 a.m.	-0.066%*	-0.556%
		TFIT03140509		
Number of events	**Date**	**Hour**	**AR**	**IVV Return**
32	**09/08/2007**	**02:45:00 p.m.**	**-0.088%***	**-0.687%**
	14/10/2008	08:30:00 a.m.	0.079%*	3.074%
	02/03/2009	**06:15:00 a.m.**	**0.093%***	**-1.979%**
		TFIT05241110		
Number of events	**Date**	**Hour**	**AR**	**IVV Return**
51	**09/08/2007**	**02:45:00 p.m.**	**-0.282%****	**-0.687%**
	29/09/2008	08:17:00 a.m.	-0.150%*	-1.444%
	14/10/2008	08:30:00 a.m.	0.250%**	3.074%
	24/01/2009	04:38:00 p.m.	0.142%*	0.470%
	01/06/2009	**07:58:00 a.m.**	**0.327%****	**0.458%**
	02/03/2009	**06:15:00 a.m.**	**0.564%*****	**-1.979%**
	25/09/2009	02:05:00 p.m.	0.441%**	0.159%
		TFIT04180511		
Number of events	**Date**	**Hour**	**AR**	**IVV Return**
50	14/10/2008	08:30:00 a.m.	0.399%**	3.074%
	23/10/2008	**08:16:00 a.m.**	**-0.175%***	**-1.932%**
	17/12/2008	**08:30:00 a.m.**	**0.286%****	**-0.914%**
	02/03/2009	**06:15:00 a.m.**	**0.354%****	**-1.979%**
	11/05/2009	02:06:00 p.m.	0.172%*	0.827%
	25/09/2009	02:05:00 p.m.	0.512%***	0.159%
		TFIT10281015		
Number of events	**Date**	**Hour**	**AR**	**IVV Return**
10	**10/09/2008**	**10:05:00 a.m.**	**-0.479%***	**0.426%**
		TFIT11241018		
Number of events	**Date**	**Hour**	**AR**	**IVV Return**
6	**02/03/2009**	**06:15:00 a.m.**	**1.467%****	**-1.979%**
		TFIT15240720		
Number of events	**Date**	**Hour**	**AR**	**IVV Return**
62	**09/08/2007**	**02:45:00 p.m.**	**-0.512%***	**-0.687%**
	25/09/2008	**09:07:00 p.m.**	**-0.591%***	**1.837%**
	14/09/2008	11:11:00 p.m.	-0.803%**	-3.268%
	29/09/2008	08:17:00 a.m.	-1.302%**	-1.444%
	03/10/2008	07:00:00 a.m.	0.548%*	1.512%
	14/10/2008	08:30:00 a.m.	2.913%***	3.074%
	28/10/2008	**11:53:00 p.m.**	**1.654%****	**0.585%**
	30/10/2008	**08:30:00 a.m.**	**1.230%****	**2.509%**
	23/10/2008	**08:16:00 a.m.**	**-1.560%****	**-1.932%**
	17/12/2008	**08:30:00 a.m.**	**0.525%***	**-0.914%**
	16/01/2009	08:02:00 a.m.	0.884%**	1.496%
	02/03/2009	**06:15:00 a.m.**	**1.112%****	**-1.979%**
	11/05/2009	02:06:00 p.m.	0.487%*	0.827%
	29/05/2009	**10:44:00 a.m.**	**0.506%***	**-0.348%**
	20/07/2009	07:36:00 p.m.	0.490%*	0.433%
	25/09/2009	02:05:00 p.m.	1.045%**	0.159%
	06/05/2010	**02:20:00 p.m.**	**-0.531%***	**1.169%**

The results in this table are obtained from matching the events from Table 5.3 to the abnormal return of a specific issue that traded in the same hour of the time of the event. The abnormal return and the significant abnormal return for a particular issue are obtained by applying the procedure described in equation (5.2) to equation (5.4). For purposes of space we just report the significant event; in the case of TFIT152240720 there were a total of 62 matches but only 17 were significant. We also report the size of the abnormal return (equation 5.3) and the contemporaneous US market proxy return (IVV return). The results highlighted in bold are negative events and all others are positive events.

Table 5.8 shows that the number of GFC events that had a significant impact on Colombian bond returns is very small compared to the number of total events. Two of the issues identified by the market model in equation (5.2) did not have any significant abnormal returns during the events of the GFC. The GFC event that had the highest impact across issues was on March 2, 2009, when AIG reported the highest loss in US corporate history. Even though the impact in the US market was negative, all significant Colombian issues reported gains on that day. One possible hypothesis would be that local investors unwound their positions abroad and recomposed their portfolios in local currencies. The negative news that had the greatest negative impact on a particular issue (TFIT15240720) was Alan Greenspan's testimony to the US Congress. The positive news with the greatest positive impact was on October 14, 2008, when all major US banks agreed to a capital injection by the Treasury. Another piece of negative news that had a negative impact on US returns, but a positive impact on Colombian bonds, was the drop in US consumer prices on December 17, 2008. An explanation could be that since the US rate is a proxy for the global interest rate, a deflation brings the expectation of new central bank expansions of liquidity and lowering interest rates, which is always good news for bond prices.

In order to test if there were statistically significant differences between abnormal returns from GFC events and those related to macroeconomic surprises, we ran a series of mean equality tests. Our hypothesis is that during the GFC, average abnormal returns for Colombian bonds should have been negative or at least lower than in other comparable periods. The same hypothesis holds for macroeconomic surprises during the GFC in order to control for confounding effects. In order to set the equality test we grouped the significant abnormal returns due to macroeconomic news surprises into three groups: 1) pre-crisis, 2) crisis; and 3) post-crisis. The pre-crisis period runs from January 9, 2007, to July 11, 2007. The crisis period runs from August 17, 2008, to December 23, 2010. The post-crisis group runs from January 19 to December 26, 2013. Finally, the GFC encompasses the dates from Table 5.7 (from August 9, 2008, to May 6, 2010). The results from the equality of means test are reported in Table 5.9.

Table 5.9: Differences in abnormal returns between GFC events and pre-crisis, crisis, and post-crisis macroeconomic news

Panel A: Average significant abnormal returns in pre-crisis period (macroeconomic surprises) and GFC significant events (in hourly percentages %)		
Pre-crisis	**GFC events**	**Difference**
-0.445%	0.288%	0.732%***
(0.0000)	(0.0001)	

Panel B: Average significant abnormal returns in crisis period (macroeconomic surprises) and GFC significant events (in hourly percentages %)		
Crisis	**GFC events**	**Difference**
0.119%	0.288%	0.168%
(0.0001)	(0.0001)	

Panel C: Average significant abnormal returns in post-crisis period (macroeconomic surprises) and GFC significant events (in hourly percentages %)		
Post-crisis	**GFC events**	**Difference**
0.008%	0.288%	-0.280%*
(0.0002)	(0.0001)	

Panel D: Average significant abnormal returns in pre-crisis period (macroeconomic surprises) and crisis (macroeconomic news) (in hourly percentages %)		
Pre-crisis	**Crisis**	**Difference**
-0.445%	0.119%	0.564%***
(0.0000)	(0.0001)	

Panel E: Average significant abnormal returns in crisis period (macroeconomic surprises) and post-crisis (macroeconomic news) (in hourly percentages %)		
Crisis	**Post-crisis**	**Difference**
0.119%	0.008%	-0.111%
(0.0001)	(0.0002)	

The results in this table report the mean equality tests for significant differences between the GFC events abnormal returns and those derived from three groups of macroeconomic surprises which are classified as: 1) pre-crisis, 2) crisis, and 3) post-crisis. The pre-crisis group of significant abnormal returns due to macroeconomic surprises comprehends the period from January 9, 2007, to July 11, 2007. The crisis group of significant abnormal returns due to macroeconomic surprises comprehends the period from August 17, 2008, to December 23, 2010. The post-crisis group of significant abnormal returns due to macroeconomic surprises comprehends the period from January 19 to December 26, 2013. We also test for differences among the three groups of abnormal returns (Panel D and E). We obtain significant abnormal returns by applying equation (5.4) $t - stat = \frac{e_{i,t}}{\sigma_{est}}$ where $e_{i,t}$ = the abnormal return and σ_{est} = the standard error of the regression. We say that an abnormal return is significant if we reject the null hypothesis of different than zero at the 90%, 95%, and 99% confidence level.

Table 5.9, Panel A, shows that the difference between the pre-crisis surprises and GFC events is significant. However, the sign of the difference is positive, which means that Colombian bonds performed better during the period of the crisis timeline than before. This is a sign of the resilience of Colombian bonds to GFC events, at least to the ones outlined in the timeline. One plausible criticism of this approach is that the transmission channel is not the GFC events but the global events. In Panel B of Table 5.9, when we test for the confounding effects due to macroeconomic surprises in the pre-crisis period and the crisis periods, the difference is not statistically different from zero. This means that the effect of macroeconomic surprises remained the same before and during the GFC. When we test for differences between GFC events and macroeconomic surprises (Panel C of Table 5.9) during the post-crisis period we find that the difference is significant. What is interesting is that the Colombian bonds performed better during the GFC period than in the period that immediately followed. Our hypothesis is that there is a significant negative effect from the European sovereign debt crisis during the post-GFC period, but that is not testable in our case due to the lack of a tradable market proxy in a high-frequency basis from Europe.

The results from Panels D and E of Table 5.9 compare macroeconomic surprises in the pre-crisis and post-crisis periods with macroeconomic surprises in the crisis period. As with the previous results, if there is no decoupling, we expect average abnormal returns of macroeconomic surprise to be negative or at least lower than those in the pre-crisis and post-crisis period. In line with GFC events, the abnormal results due to surprises in the crisis period are positive and higher than in the pre-crisis period (Panel D), and in the post-crisis period the difference in abnormal returns is not different from zero. Therefore, there is evidence of decoupling in the case of Colombian currency bonds from the events identified by the GFC timeline.

The results obtained in this paper corroborate at the intraday level the decoupling of emerging markets during the GFC as suggested by studies done at the aggregate level (Dooley and Hutchison, 2009). It also corroborates, just at the specific case level, another aggregate study by Dungey et al. (2010), which found that the US subprime crisis had only a small impact on the volatility of emerging-market sovereign bond market returns.

5.5 Conclusion

We use high-frequency hourly data to analyse the effects of the events that surrounded the GFC on Colombian bond prices. Our study is one of the few event studies that analyse the effect of the GFC on emerging-market bond returns. In order to control for possible confounding effects surrounding the events of the GFC, we analyse the effect of macroeconomic surprises on Colombian bond prices before, during, and after the crisis. We find that for all the periods under observation local news related to inflation had the greatest impact in bond prices. In the case of global and regional news, inflation- and trade-related surprises also had significant effects on bond prices, but to a lesser extent. Our results show that there was resilience (in terms of abnormal returns) in Colombian bond returns to events derived from the GFC. We also find that, on an average, Colombian bonds performed better during the period of the GFC than during the periods before and after the GFC. This finding is further reinforced by the fact that after controlling for confounding effects such as other regional and local macroeconomic announcements abnormal results due to surprises in the crisis period are positive and higher than in the pre-crisis period.

Finally, our conclusion is that, at least for the events described in the GFC timeline, there is evidence of decoupling in the case of Colombian local-currency bonds.

Chapter 6 Conclusion

The general conclusion for this thesis is that there is evidence that there was resilience or decoupling in the case of government-sponsored or regulated assets during certain phases of the Global Financial Crisis (GFC) at least for the assets under study. In the case of the Colombian pension funds (Chapter 2), using a time-varying framework we were able to analyse the behaviour of a particularly isolated (autarchic) asset during the GFC. Furthermore, our time-varying framework allows us to decompose risk into its systematic and idiosyncratic components by assuming that all shocks are transmitted through a common factor, which in our case was the S&P 500 index and key bond indices in order for further test contagion by asset type in the case of the Colombian private pension funds. Using this decomposition approach, we were able to test for contagion in the context of its two broader definitions: 1) As creating additional linkages in market co-movement; 2) As creating additional volatility spillover from one market to the other. Although the proposed models for testing contagion are commonly used in the literature, our contribution is that by allowing the models to capture the effect of leverage, new insights can be gained into the nature and the magnitude of contagion in autarchic asset returns. Additionally, by using quantile autoregression to test the significance of crisis dates using time-varying systemic risk, it is possible to overcome some of the limitations and biases generated by asymmetric data. Our results showed that in the case of the Colombian funds there was no evidence of contagion during the first two episodes of the GFC when we allowed for time-varying volatilities and different common factors. However, during the ESD crisis episode there is strong evidence of contagion in all tests, including robustness checks.

In Chapter 3 we estimate a four-factor quantile regression model of returns to BRIC-economy SOEs; we study their performance during the financial crisis and recovery period. Our model benchmarks the SOEs against US market factors, including industry-sector indexes. US market, size, and value-growth factors are significant factors for excess returns for SOEs from many countries and sectors. Additionally, by analysing dependence at extreme quantiles during crisis episodes, we observe that indeed certain SOE industry sectors were either more protected from or more exposed to contagion than others. Our results show that there is evidence of a “cushion effect” against contagion in the telecommunication, utilities, materials, information technology, and energy SOE sectors during different episodes of the crisis. Our results are similar to those obtained in a recent study in the area of segmentation that shows there is a trend of reversal in global market integration at the industry and country level (see: Bekaert, Harvey, et al., 2011). Particularly worth noting is the increase in segmentation in the industrial sectors of emerging markets during crisis periods. Finally, this chapter has important implications for the literature concerning the role of government ownership and intervention by addressing the issue of performance of partially privatised SOEs, and stating that this “partnership” between the state and private investors can be mutually beneficial during crisis periods, especially due to the “cushion effect” or decoupling in the tails when extreme negative movements impact on certain economic sectors.

In Chapter 4, by using a factor model based on country-specific and common market determinants of sovereign spreads in the context of propensity-matching estimators, we propose a novel framework to test for differences in sovereign debt spreads. Our findings suggest that the most common country-specific factor among portfolio groups is the local equity premium, with the exception of the Eurozone and

PIIGS, where changes in neighbouring countries (regional portfolio) have a larger effect. However, in the specific case of the GFC there were different channels for transmission of contagion. The most common channel of contagion transmission among most portfolio groups was macroeconomic fundamentals related to liquidity or "wake-up" contagion, where investors pay close attention to the country's ability to meet its financial obligations. There was evidence of a latent-factor or "pure" contagion in all of the portfolios, with the exception of the all-countries portfolio, suggesting that most of the latent factors in the remaining portfolios can be explained by changes in the global bond portfolio.

In analysing the behaviour of international government bonds during the crisis, the evidence shows that the most meaningful period for differences in sovereign spreads was the ESD phase, in which the spreads rose substantially from the previous crisis periods. Additionally, in our proposed framework we define our test for differences in spreads as a statistically significant change in the average change in spreads between the observations in the counterfactual non-crisis period and those of the crisis period. We do this in order to determine the actual economic significance in basis points of the different phases of the crisis versus non-crisis periods. In order to do this we define this average change as our average treatment effect on the treated (ATET), where the "treatment" is the crisis period. In this way, we are able to obtain estimates based on a similar distribution between the crisis and counterfactual group and reduce the problem of selection bias. We compare the ATET results with tests using more conventional sampling: both equal, unequal, and overlapping non-crisis periods, and also using different kernel specification of matching estimators without significant changes in our main results. Using our preferred method, which is robust to endogeneity, we found evidence that the portfolio of local-currency emerging-market debt did not exhibit any significant difference in spreads during the GFC. This means that based on the common characteristics of the counterfactuals, the emerging countries that issued debt in local currency have dealt with similar economic conditions in the past and were decoupled from the effects of the GFC.

Finally, in Chapter 5 we further corroborate the findings of Chapter 4 by analysing the effects of the events of the GFC on Colombian local-currency bonds using high-frequency data. Additionally, in order to control for possible confounding effects surrounding the events of the GFC, we analyse the effect of macroeconomic surprises on Colombian bond prices before, during, and after the crisis. We find that for all the periods under observation, local news related to inflation had the greatest impact on bond prices. In the case of global and regional news, inflation- and trade-related surprises also had significant effects on bond prices, but to a lesser extent. Our results show that there was resilience/decoupling (in terms of abnormal returns) in Colombian bond returns to events derived from the GFC. We also find that, on average, Colombian bonds performed better during the period of the GFC than the period before and after the GFC.

6.1 Future research

We hope that by analysing the effect of government regulation in emerging markets there can be some lessons to be learned in regard to what special kinds of assets can help to mitigate the effects of contagion and the consequent loss of wealth, which can have devastating effects in emerging economies. In the case of Chapter 2, future research in this area should take into account the role of regulatory constraints on portfolio holdings. Even though these “quantitative restrictions” can limit the risk/return potential of the autarchic portfolio, these same restrictions can limit potential losses in times of crisis. In the case of Chapter 3, future research could take into account the different levels of government ownership and indirect intervention (subsidised loans, effect of government contracts, etc.) as well as the effects that "implicit" government guarantees have on firm performance in both developed and emerging markets. In the case of Chapter 4, the proposed contagion-testing framework can be applied to other financial assets under different kinds of setups to measure the effect of different government regulations on asset prices. Finally, the method proposed in Chapter 5 can be used as a common framework for analysis of the effects of the GFC using other sources of high-frequency data of financial assets in other emerging markets.

Appendix A

According to the guidelines set in the *Decreto 1592 de 2004,* Articles 1 and 2, the formula is:

$$RM = 0.5(0.7RPS)+0.5(w_1*0.7RRV+ w_2*0.7RFE+ w_3*0.7BM)$$

Where:

RM = Minimum guaranteed annual rate of return.

RPS = Weighted average rate of return of all the private pension funds that compose the system per annum.

RRV = Annual rate of return of the Colombian stock market index

RFE = Annual rate of return of a global stock market chosen by the Colombian banking regulatory agency.

BM = Annual return of a benchmark portfolio chosen and calculated by the Colombian banking regulatory agency.

w_1 = The proportion of the assets of the fund invested in Colombian stocks.
w_2 = The proportion of the assets of the fund invested in foreign stocks.
w_3=The proportion of the assets of the fund invested in other assets.

Appendix B

According to the guidelines set in the Chapter 12 of the *Circular Externa 036 de 2003,* Article 1, pages 1 to 4, the value of a PPF and how is expressed in unit value should be reported daily and in unit values. The units should reflect the market value of the contributions of the affiliates and their number represents the equity position of the affiliate in the fund. The daily variation in the unit value should reflect the return between the buy and sell dates (Superfinanciera, 2003).

The formulas for calculating the value of the fund and the values of the units are:

For the value of the fund:

VFC = VFI + AT - TR - RC- MP - RAV - DS - SP - CA - OC - FSPsl - FSPsbp - FSPsbaf -FGPM - RV - OR ± AN

Where:

VFC = Value of the fund at the closing of day (t) – before dividends and yields.

VFI = Value of the fund at the opening of day (t).

AT = Contributions and transfers received in day (t). It includes deposits for severance fund subsidies. In the case of mandatory pension funds in the brute value of the deposit is all inclusive (it includes commissions, insurance, commissions, and all relevant subsidies or additional contributions set by the legislation).

TR = Value of the transfers to other private pension fund administrator by petition of the client at day (t).

RC = Value of payments of severance fund subsidies (partial or definitive) at day (t).

MP = Value of the pension allowances paid at day (t).

RAV = Value of payments related to cancellations of voluntary contribution plans at day (t).

DS = Value of reversals in outstanding balances to affiliates or pensioners of the private pension fund in day (t).

SP = Value of pension insurance payed at day (t).

CA = Value of the commission payable to the private pension fund administrator at day (t).

OC = Value of other commissions payable to the private pension fund administrator at day (t).

FSPsl = Value of the contributions to the solidarity fund, subaccount solidarity, at day (t).

FSPsbp = Value of the contributions to the solidarity fund, subaccount subsistence pensioners, at day (t).

FSPsbaf = Value of the contributions to the solidarity fund, subaccount affiliates, at day (t).

FGPM = Value of the contributions of the minimum guarantee pension fund at day (t).

RV = Value of balance transfers to insurance companies for pension payments under the perpetual rent scheme at day (t).

OR = Value other transfers/retirement at day (t).

AN = Value of cancelled operations at day (t).

For the number of units:

NUC = NUCI + NUAT - NUTR - NURC – NUMP - NURAV - NUDS - NUSP - NUCA - NUOC - NUFSPsl - NUFSPsbp - NUFSPsbaf - NUFGPM - NURV - NUOR ± NUAN

Where:

NUC = Number of units of the fund at the closing of day (t), which must be the same number at the opening of the next day (t+1).

NUCI = Number of units of the fund at the opening of day (t).

NUAT = Number of units related to new contribution and incoming transfers at day (t).

NUTR = Number of units related to transfers of existing contributions to other private pension fund administrators at day (t).

NURC = Number of units of severance fund subsidies (partial or definitive) at day (t).

NUMP = Number of units related to the payment of pensions allowances payed at day (t).

NURAV = Number of units corresponding to the value of payments related to cancellations of voluntary contribution plans at day (t).

NUDS = Number of units corresponding to the value of reversals in outstanding balances to affiliates or pensioners of the private pension fund in day (t).

NUSP = Number of units corresponding to the value of pension insurance payed at day (t).

NUCA = Number of units corresponding to the value of the commission payable to the private pension fund administrator at day (t).

NUOC = Number of units corresponding to the value of other commissions payable to the private pension fund administrator at day (t).

NUFSPsl = Number of units corresponding to the value of the contributions to the solidarity fund, subaccount solidarity, at day (t).

NUFSPsbp = Number of units corresponding to the value of the contributions to the solidarity fund, subaccount subsistence pensioners, at day (t).

NUFSPsbaf = Number of units corresponding to the value of the contributions to the solidarity fund, subaccount affiliates, at day (t).

NUFGPM = Number of units corresponding to the value of the contributions of the minimum guarantee pension fund at day (t).

NURV = Number of units corresponding to the value of balance transfers to insurance companies for pension payments under the perpetual rent scheme at day (t).

NUOR = Number of units corresponding to the value of other transfers/retirement at day (t).

NUAN = Number of units corresponding to the value of cancelled operations at day (t).

For all purposes, when a new fund enters the market the initial net asset value (NAV) is set at COP$10,000. The total value of the funds including dividends and coupon payments is equal to:

VFCR = VFC + INGt - GTSt

Where:

VFCR = Value of the fund at the closing of day (t) including yields from related investments which must be the same as the opening of the next day (t+1).

VFC = Value of the fund at the closing of day (t) – before dividends and yields.

INGt = Yields from investments in the fund at day (t).

GTSt = Expenses related to purchasing investments for the fund at day (t).

Finally the net asset value of the unit (VUC) at the end of day (t) is equal to VUC=VFCR/NUC.

Appendix C

Quantile autoregression as proposed by Koenker and Xiao (2006) can be an extremely useful econometric tool when dealing with asymmetric series. Using Koenker and Xiao (2006) notation, assume that (U_t) is a normal distributed variable that exhibits the following (n) autoregressive process where the parameters $\theta_{n's}$ are unknowns:

$$y_t = \theta_0(U_t) + \theta_1(U_t)y_{t-1} + \theta_2(U_t)y_{t-2} + ... + \theta_n(U_t)y_{t-n} \qquad (1)$$

Furthermore, if y_t is a monotone increasing function of (U_t) and that τ is a conditional quantile function of y_t then equation (1) can be rewritten as:

$$Q_{y_t}(\tau | y_{t-1}...y_{t-n}) = \theta_0(\tau) + \theta_1(\tau)y_{t-1} + \theta_2(\tau)y_{t-2} + ... + \theta_n(\tau)y_{t-n} \qquad (2)$$

In matrix form:

$$Q_{y_t}(\tau | F_{t-1}) = \mathrm{X}_t^T\theta_0(\tau) \qquad (3)$$

Where $\mathrm{X}_t^T = (1, y_{t-1}, ...y_{t-n})^T$ and F_{t-1} is the sigma field generated by $\{y_s, s \le t\}$. The intuition behind the quantile autoregression model is the fact that the autoregressive coefficients or (θ) are quantile dependent (τ) thus allowing for the coefficients to vary along the quantiles. Therefore a QAR(1) process can be defined as:

$$Q_{y_t}(\tau | F_{t-1}) = \theta_0(\tau) + \theta_1(\tau)_{y_{t-1}} \qquad (4)$$

Where $\theta_0(\tau) = \sigma\phi^{-1}(\tau)$ and $\theta_1(\tau) = \min\{\gamma_0 + \gamma_1\tau, 1\}$ for $\gamma_0 \in (0,1)$ and γ_1 >0. According to Koenker and Xiao (2006) if $U_t > (1-\gamma_0)/\gamma_1$ [42] then the model generates a non-stationary process, but for smaller values of U_t there is a mean reversion tendency. This means that even though the whole process of of y_t can be globally stationary, it can still display asymmetric persistence in the presence of large shocks. Therefore, one of the interesting qualities of a QAR process is that it allows for some form of asymmetric behavior in some quantiles while maintaining stationarity in the long run.

[42] This come from the fact that $Q_{g(U)^{(\tau)}} = g(Q_u(\tau)) = g(\tau)$ therefore U is a function of the quantile (see: Koenker and Xiao (2006, p. 981)

Appendix D

In order to check the robustness of our results, we use two other algorithms based on kernel matching. The difference between neighbour matching and kernel algorithms is that the former assign equal weights to all matched observations drawn from the non-crisis period and the latter give more weight to the observations that are more closely matched. Following the implementation procedure used by Nssah (2006), the proportional weight assigned to the observations in the non-crisis period (p_j) as a function of how close they are to the crisis period (p_i) is:

$$w_{ij} = \frac{K\left(\frac{p_i - p_j}{h}\right)}{\sum_{j\in\{d=0\}} K\left(\frac{p_i - p_j}{h}\right)} \quad , \tag{3.14}$$

where w_{ij} is the assigned weight, and h is the bandwidth, which is set at a fixed value of 0.06 (Becker and Ichino, 2002). We define K as the Gaussian (GAUSS) and Epanechnikov kernel (EPNK):

$$\begin{aligned} K(u) &= \frac{1}{\sqrt{2\pi}}\exp\left(\frac{-u^2}{2}\right) \\ K(u) &= \frac{3}{4}(1-u^2)\,\mathrm{xI}\left(|u| \le 1\right) \end{aligned} \tag{3.15}$$

where $u=(p_i - p_j)/h$ and I is an indicator variable that takes the value of 1 (true) and 0 (false) when the condition $\|p_i - p_j\| < 0.06$ is met. For these two kernels, we also establish an area of common support based on the minimum and maximum propensity scores in the crisis period obtained from equation (3.66). This limits the sample observations drawn from the non-crisis period to those values within the range of those in the crisis period, further reducing the possibility of biases due to outliers.

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