The Curse of Non-Investment Grade Countries: Excess Vulnerability.[1]

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Abstract

Mexico was upgrade from non-investment to investment grade in March of 2000. This paper examines the impact of this event on the stochastic properties of the transmission of shocks between Argentina and Mexico. The paper shows that there is a statistically significant change in the propagation of shocks the day the upgrade was announced. Furthermore, it is found that the parameters that shifted are those explaining the diffusion of shocks through the means, while the transmission through the variances remained stable. From the methodological point of view, the paper offers an identification procedure based on conditional heteroskedasticity (ARCH) that solves the problem of estimation in a simultaneous equations model.

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1 Introduction

On March 7th, 2000, Moody's upgraded Mexican Debt to investment grade.[1] Obviously, this improved the external and internal financing conditions for Mexico: average yields, and their conditional variances came down almost immediately. The benefits from the credit rating upgrade were not surprising.[2] Nevertheless, what has been striking is the fact that since the upgrading Mexico seemed immune to the crisis in Turkey and Argentina.

Indeed, Argentinean and Mexican bond yields used to move closely in the past. However, after the upgrade they look as if they are independent. Figure 1 shows the daily yields on Argentinean and Mexican sovereign bonds from January 1997 to May 2001. The data corresponds to the stripped yield on the JPMorgan's EMBI+ indexes of these two countries. A casual observation of the data suggests that the two markets co-moved strongly prior to March 2000, but have drifted apart since then. The same behavior can be found in the conditional volatilities. In Figure 2, the rolling variance of the yields using a 30 days window is depicted. Observe that the variances were almost aligned prior to March 2000[3]. During the Russian Cold (August 1998) the volatility of both yields increased by 10 times, and around the Brazilian Sneeze (January 1999) the variance raised significantly too. Nonetheless, throughout the recent Argentinean turmoil (the Pampa Shivering) the variance in Mexico has been unaffected, while the Argentinean variance has increased in more than 5 times. These two pieces of evidence hint that Argentina and Mexico are co-moving less today than years before. Therefore, it is not shocking that the correlation coefficients have dropped. In Figure 3, the rolling conditional correlation of the yields is portrayed. The correlation was extremely high during the Russian Cold, mainly as the outcome of a large common shock, but it came down to close to 70 percent during 1999-2000. After March of 2000, it has decreased to less than 50 percent.

The objective of this paper is to study, first, whether or not the transmission of shocks was significantly altered by the rating upgrade. Second, it analyzes in which dimensions the propagation mechanism between Argentinean and Mexican sovereign bonds has changed. Third, it evaluates the predicted fall in the correlations that can be attributed to the event.

Before summarizing the results in the paper, it is worth asking why this change in credit rating is of particular interest. Certainly, upgrades and downgrades have occurred previously. However,

[1] S&P upgraded Mexico (on March 13th), but not all the way to investment grade.

[2] Kaminsky and Schmukler [2001] offer evidence on how upgrades and downgrades affect country risk and stock returns.

[3] This fact is consistent with Edwards [1998] and Edwards and Susmel [2000] where strong variance spillovers are found.

in the last five years in the Latin American region, there has been only two cases where the rating has moved between investment and non-investment grades. In 1999, Colombia was downgraded from investment grade to non investment grade, and in 2000, Mexico was upgraded. Because of regulatory restrictions, these two changes imply a shift in the investor universe. For example, in the case of Mexico, the upgrade means that now a broader set of investors, such as insurance companies, pension funds, and certain mutual funds, can hold Mexican debt. Thus, theories of contagion based on the identity of the investors can be tested in this experiment. The other changes in rating do not modify the type of investor holding the instruments. Hence, they are not able to assess the significance of segmentation in the investors type as the source of co-movement.

The three main results of the paper are: First, it is possible to reject the hypothesis that the propagation of shocks is stable after March 7 of 2000. It is important to highlight that the evidence provided in Figures 1 to 3 is at best suggestive. In order to test for the stability of parameters, a more powerful and robust test is implemented: the DCC test. This procedure allows to test for parameter stability of a linear multivariate model in the presence of simultaneous equations, omitted variable biases, and conditional heteroskedasticity, which is the set up studied here. The stability in the relationship is covered in Section 2.

The second result is related to how the propagation has varied. In particular, the transmission mechanism is split in two channels: the spread of shocks that occurs through the means, and the diffusion that takes place through the second moments. Unfortunately, the DCC only provides evidence that a regime shift occurred, but does not indicate which parameters have changed. This procedure is able to disentangle the diffusion of shocks that takes place through the means, and through the second moments. The contagion literature mostly studies the propagation by the means. There are three notable exceptions: Dungey and Martin [2001], Edwards [1998], and Edwards and Susmel [2000]. These papers study the propagation of shocks in ARCH or GARCH specifications. In order to deal with the simultaneous equations issue, however, they assume exclusion restrictions on the ARCH structure (Dungey and Martin) or estimate the model on the reduced form (the two Edwards' papers). The present paper extends this literature and offers a new procedure that is able to solve the simultaneous equations problem with a weaker set of assumptions. Section 3 estimates the contemporaneous propagation between Argentinean and Mexican sovereign bonds using an identification procedure based on conditional heteroskedasticity. The main conclusion from this exercise is that there is a sizeable reduction on the propagation of shocks through the means, while the diffusion via second moments has remained relatively stable.

The third result evaluates the predicted reduction in the co-movement as the consequence of the change in parameters. Section 4 shows that the change in parameters imply a reduction in the predicted unconditional correlation from 80 percent to 50 percent. In other words, around a third of the co-movement during the non-investment grade period was explained by the segmented market. This finding supports those theories of contagion in which the propagation of shocks is due to the type of investors. Indeed, it can be argued that the only change between Mexico and Argentina in March 2000 is the fact that now a larger, and different, set of investors is able to hold Mexican bonds. The paper shows that a sizeable proportion of the co-movement is explained by this channel. Finally, conclusions are presented in Section 5.

2 Is the propagation across countries stable?

Confirming Figures 1 to 3, Table 1 shows the reduction in correlation between other major Latin American Countries and Mexican bonds. These are simple correlations, and therefore, an incomplete picture of the actual shift. Nevertheless, it is striking that all correlation coefficients dropped significantly after March 7th of 2000.[4]

Mexico Correlation with	Arg	Bra	Col	Per	Ven	EMBI+
March 8, 1998 - March 6, 2000	88.3%	95.3%	76.3%	91.6%	79.1%	97.0%
March 8, 1999 - March 6, 2000	88.3%	91.1%	87.5%	95.0%	89.1%	93.9%
March 7, 2000 - May 28, 2001	42.4%	66.3%	33.6%	35.1%	74.9%	73.3%

Table 1: Conditional simple correlations.

These stylized facts are not entirely conclusive about the stability of parameters. As has been argued by Ronn [1998] the use of conditional correlations to assess regime shifts could produce misleading conclusions.[5] In order to test for a regime change, then, I run the DCC test.[6] The DCC is designed to test for the stability of a set of multinomial variables that are simultaneously determined, have omitted variable problems and suffer from heteroskedasticity. Previous tests have been unable to handle all three issues at the same time.

The test is based on the assumption that the data can be divided in two sub-samples with a known break. The null hypothesis is that the heteroskedasticity is explained by the shift in the variance of only one of the shocks in the original system. The alternative hypothesis is that the

[4] In the case of Colombia, the sample only includes the period after Colombia was downgraded to non-investment grade.

[5] For the limitations on the correlation coefficient see: Ronn [1998], and its applications to contagion: Boyer, Gibson and Loretan [1999], Forbes and Rigobon [1998], and Loretan and English [2000].

[6] DCC stands for *Determinant of the Chance in Covariance* matrices. See Rigobon [2000c] for a detailled description of the test. For application see Killeen, Lyons and Moore [2001]. They test the stability of the relationship between European exchange rates after the announcement of joining EMU. Additionally, see Rigobon [2000b] for an application on the stability of the transmission of shocks across stock markets.

change in the volatilities is explained either by parameter instability or by all variances shifting. The test is implemented by (i) computing the covariance matrix in each of the sub-samples, (ii) subtracting the estimated matrices, and (iii) calculating the determinant of the difference. Under the null, the determinant should be equal to zero. The drawback of the DCC test is that if it is rejected, it does not provide guidance on the reason of the rejection. In this section, I concentrate on the first question, and once the stability hypothesis has been rejected, the next section studies in which dimensions the relationship has varied.

The first step to implement the test is to define the date of the break. In the Mexican case, the break was chosen to be March 7th of 2000.

The second step is to delineate the two sub-samples where the test will be performed. In other words, the test requires the determination of the windows where the covariance matrices will be computed and compared. The choice of the length of the windows is not innocuous. On the one hand, the test is rejected if the heteroskedasticity is explained by changes in the volatilities of all structural shocks. Which is likely to occur if the windows are too large. From the parameter stability point of view, a rejection because the variance of all structural shocks is changing is non-interesting. Therefore, the importance of defining the windows narrowly enough. On the other hand, if the windows are too small, then the estimates of the covariance matrix are noisy and the size of the test is unreliable. Hence, I use three different windows: For the non-investment grade period I use the 30 days prior to the upgrading, and for the investment grade period I use 10 or 20 or 30 days.

The third step in the DCC is to compute the covariance matrices, and calculate the determinant of their difference. In Table 2, the results are presented. The asymptotic distribution of the test has not been studied, and most of the results are based on bootstraps. 1000 draws are computed and it has been assumed that the covariance matrices are correlated across the two samples. The first row corresponds to the mean of the bootstrapped distribution, the second row is its standard deviation, and the last row is the mass above zero. The distributions are in general not normal. Thus, it is common that the standard deviations are large relative to their means. Thus, most of the analysis is based on the mass above zero: in other words, if the proportion of realizations of the bootstrap that have positive determinant is small, then the null hypothesis is rejected.

	Argentina versus Mexico		
Determinant	10 days	20 days	30 days
Mean of Distribution	-27.28	-38.80	-25.73
Standard Deviation	26.67	23.28	15.09
Mass above Zero	6.8%	0.5%	1.3%

Table 2: DCC Test. Argentina versus Mexico.

As can be seen, the mass above zero is small for all three windows. The proportion of realization with positive determinant are 6.8, 0.5 and 1.3 percent when the windows are 10, 20, and 30 days, respectively. These results imply a strong rejection of the null hypothesis even at short horizons. Hence, it is safe to conclude that the transmission mechanism between Mexico and Argentina was altered by the rating upgraded.

This result supports the theories of contagion in which the identity of investors explains a sizeable proportion of the co-movement. As was mentioned before, in the last 5 years only two changes in credit ratings have implied shifts in the investors universe: Colombia in 1999, and the case studied in this paper. However, a multitude of other downgrades and upgrades have occurred. Nevertheless, those changes have not produced a significant shift in the propagation mechanism. It is important to highlight, then, that the shift in the propagation of shocks is due to the modification of the type of investor, and not to a liquidity shock. In the sample analyzed here, there were no margin calls nor liquidity shocks in the bond market. Therefore, the drop in co-movement is not due to the disappearance of a liquidity shock, but by the shift in the investor universe.

Indeed, in the Latin American sovereign bond market the three instances in which the hypothesis of stability has been rejected are: January-March 1995, August 1998, and March 2000. The first two cases, involved large liquidity shocks to market participants. The second one implies a swing in the investor set. Moreover, the first two shocks have been short lived, while this one seems to be long lasting.[7] There have been very few theories that explain the co-movement across markets based on the identity of investors, though. Most of the papers study liquidity shocks in a segmented market, but few study the co-movement driven by the segmentation of the market by itself. Two notable exceptions are Kaminsky and Reinhart [1998] and Gromb and Vayanos [2001]. Kaminsky and Reinhart study the increase in correlation due to common lenders, while Gromb and Vayanos developed a model of asset price co-movement in segmented markets with wealth effects. This evidence supports the importance of these channels in the sovereign bond market.

3 Estimating the propagation mechanism between Argentina and Mexico

The previous section shows that Mexican's upgrade altered the diffusion of shocks between Argentina and Mexico. Unfortunately, the DCC is unable to indicate what coefficients, or aspects of the relationship, have changed. In this section, I explore this question. In particular, I analyze

[7] See Rigobon [2001a].

how the diffusion through means and second moments was affected. In order to do so, I propose a "structural" ARCH model approach.

3.1 Identification in a structural ARCH model.

This section derives the structural ARCH approach, and studies the conditions for its identification and estimation. To clarify the intuition, the problem is discussed in the general framework of simultaneous equations. Consider the following system of equations:

$$\begin{aligned} Arg_t &= \beta Mex_t + \varepsilon_t, \qquad (1) \\ Mex_t &= \alpha Arg_t + \eta_t, \qquad (2) \end{aligned}$$

where (1) is the Argentinean equation, (2) is the Mexican equation, Arg_t is the observed yield in Argentinean bonds, Mex_t is the observed yield in Mexican bonds, and ε_t and η_t are the structural shocks. Additionally, assume the structural shocks satisfy the following ARCH model:

$$\begin{aligned} \varepsilon_t &= \sqrt{h_{\varepsilon,t}} \cdot v_{\varepsilon,t} \qquad (3) \\ \eta_t &= \sqrt{h_{\eta,t}} \cdot v_{\eta,t} \qquad (4) \end{aligned}$$

where

$$\begin{aligned} E\left(v_{\varepsilon,t}\right) &= 0 \quad E\left(v_{\varepsilon,t}^2\right) = 1, \\ E\left(v_{\eta,t}\right) &= 0 \quad E\left(v_{\eta,t}^2\right) = 1, \\ E\left(v_{\varepsilon,t} v_{\eta,t}\right) &= 0. \end{aligned}$$

Furthermore, assume that the conditional variances satisfy

$$\begin{pmatrix} h_{\varepsilon,t} \\ h_{\eta,t} \end{pmatrix} = \begin{pmatrix} \zeta_\varepsilon \\ \zeta_\eta \end{pmatrix} + \lambda \begin{pmatrix} \varepsilon_{t-1}^2 \\ \eta_{t-1}^2 \end{pmatrix}, \qquad (5)$$

where ζ_ε and ζ_η are positive constants, and

$$\lambda \equiv \begin{bmatrix} \lambda_{\varepsilon\varepsilon} & \lambda_{\varepsilon\eta} \\ \lambda_{\eta\varepsilon} & \lambda_{\eta\eta} \end{bmatrix}. \qquad (6)$$

Equations (1) to (6) is what is defined as the structural model.

The most important assumption that has been imposed in the model is the zero-correlation between the structural shocks: $E(\nu_{\varepsilon,t}\nu_{\eta,t}) = 0$, which implies that $E(\varepsilon_t\eta_t) = 0$. Actually, this covariance restriction, together with the existence of heteroskedasticity, are the identifying restrictions. Here, I explore the rationale behind the assumption, while leaving the discussion of the necessity of the restriction for the Appendix (See Appendix B).

The data analyzed corresponds to the stripped yield of Argentinean and Mexican foreign currency denominated debt. Therefore, by construction, this data has no currency risk, nor US interest rate risk (these are the spreads over US interest rates). Innovations to these series are mainly default risk, or country risk. It is reasonable to assume, then, that if the countries were completely isolated, the innovations to default risk in Mexico should be uncorrelated to the innovations in Argentinean default risk. In other words, under autarky ($\alpha = \beta = 0$) the stripped yields should be uncorrelated. This structural interpretation of the equations grants the covariance restriction.

In this model, the only interrelationship across the country risks comes from the fact that the two countries trade, share macroeconomic policies, export similar goods to third markets, have common lenders, or share investors. These channels are summarized by the contemporaneous coefficients: α, β, and λ. α and β reflect the diffusion that occurs by the means, while the matrix λ summarizes the spread of shocks through the second moments.

There are three minor remarks that deserve some attention. First, the structural model assumes the standard supply and demand setup. This is equivalent to assume a latent factor model too. Both models have the exact same problems of identification, and essentially, one is a transformation of the other. See Appendix C for a detailed derivation. Second, notice that it has been assumed that the ARCH effects only include one lag. This assumption is harmless in terms of the properties of the identification problem discussed here. Third, λ is not triangular. Observe that if (at least) one of the off-diagonal elements of λ is zero, then the system is identified. This is equivalent to an exclusion restriction. See Dungey and Martin [2001] for the estimation of multivariate GARCH models in which this assumption has been imposed. However, in this application, there is no reason to assume that the volatility in one market has no effect on the future variance of the other one.

It is well known that equations (1) and (2) cannot be estimated consistently with standard procedures. However, in terms of the standard ARCH and GARCH literature, the simultaneous equation issue does not represent a problem. Most of the literature studies the predictability of the variables. Therefore, the estimation is almost always implemented on a reduced form. In this paper, I look at a different problem. I am interested in identifying a structural form by using the reduced form estimates.

In Appendix A the reduced form is derived from the structural equations. Here, I summarize the relevant equations. The reduced form residuals are given by:

$$\omega_{Arg,t} = \frac{1}{1-\alpha\beta}(\beta\eta_t + \varepsilon_t) \tag{7}$$

$$\omega_{Mex,t} = \frac{1}{1-\alpha\beta}(\eta_t + \alpha\varepsilon_t) \tag{8}$$

where $\omega_{Arg,t}$ and $\omega_{Mex,t}$ have mean zero. The conditional moments are given by $\Sigma_{\omega,t}$

$$\Sigma_{\omega,t} = \begin{bmatrix} \omega^2_{Arg,t} & \omega_{Arg,t}\omega_{Mex,t} \\ & \omega^2_{Mex,t} \end{bmatrix}$$

The covariance of the reduced form residuals is different from zero because of the simultaneous equation parameters are different from zero too. Denote the elements of the expected conditional covariance matrix of the reduced form as

$$E\Sigma_{\omega,t} = \begin{bmatrix} h_{Arg,t} & h_{ArgMex,t} \\ & h_{Mex,t} \end{bmatrix}.$$

where they satisfy the following ARCH model

$$\begin{bmatrix} h_{Arg,t} \\ h_{ArgMex,t} \\ h_{Mex,t} \end{bmatrix} = \begin{bmatrix} \hat{\zeta}_{Arg} \\ \hat{\zeta}_{ArgMex} \\ \hat{\zeta}_{Mex} \end{bmatrix} + \frac{1}{1-(\alpha\beta)^2} A \begin{bmatrix} \omega^2_{Arg,t-1} \\ \omega^2_{Mex,t-1} \end{bmatrix}, \tag{9}$$

where $\hat{\zeta}_{Arg}$, $\hat{\zeta}_{ArgMex}$, and $\hat{\zeta}_{Mex}$ are constants given by

$$\begin{bmatrix} \hat{\zeta}_{Arg} \\ \hat{\zeta}_{ArgMex} \\ \hat{\zeta}_{Mex} \end{bmatrix} = \frac{1}{(1-\alpha\beta)^2} \begin{bmatrix} \beta^2 & 1 \\ \beta & \alpha \\ 1 & \alpha^2 \end{bmatrix} \begin{bmatrix} \zeta_\eta \\ \zeta_\varepsilon \end{bmatrix}, \tag{10}$$

and where A is a three by two matrix given by

$$A \equiv \begin{bmatrix} \left[\beta^2\lambda_{\eta\varepsilon} + \lambda_{\varepsilon\varepsilon}\right] - \alpha^2\left[\beta^2\lambda_{\eta\eta} + \lambda_{\varepsilon\eta}\right] & -\beta^2\left[\beta^2\lambda_{\eta\varepsilon} + \lambda_{\varepsilon\varepsilon}\right] + \left[\beta^2\lambda_{\eta\eta} + \lambda_{\varepsilon\eta}\right] \\ \left[\beta\lambda_{\eta\varepsilon} + \alpha\lambda_{\varepsilon\varepsilon}\right] - \alpha^2\left[\beta\lambda_{\eta\eta} + \alpha\lambda_{\varepsilon\eta}\right] & -\beta^2\left[\beta\lambda_{\eta\varepsilon} + \alpha\lambda_{\varepsilon\varepsilon}\right] + \left[\beta\lambda_{\eta\eta} + \alpha\lambda_{\varepsilon\eta}\right] \\ \left[\lambda_{\eta\varepsilon} + \alpha^2\lambda_{\varepsilon\varepsilon}\right] - \alpha^2\left[\lambda_{\eta\eta} + \alpha^2\lambda_{\varepsilon\eta}\right] & -\beta^2\left[\lambda_{\eta\varepsilon} + \alpha^2\lambda_{\varepsilon\varepsilon}\right] + \left[\lambda_{\eta\eta} + \alpha^2\lambda_{\varepsilon\eta}\right] \end{bmatrix}. \tag{11}$$

The reduced form is defined by equations (7) to (11). This is a restricted multivariate ARCH that can be estimated by GMM or simulated moments. Furthermore, note that the structural parameters are identified from the reduced form estimates. From the multivariate ARCH, A can be estimated, providing six equations that have to be explained by six structural coefficients.

Observe that the ARCH model to be estimated is quite different from the typical multivariate setup in the literature. In fact, the standard multivariate ARCH has the following structure:

$$\begin{bmatrix} h_{Arg,t} & h_{ArgMex,t} \\ & h_{Mex,t} \end{bmatrix} = \begin{bmatrix} \hat{\zeta}_{Arg} & \hat{\zeta}_{ArgMex} \\ & \hat{\zeta}_{Mex} \end{bmatrix} + B' \begin{bmatrix} \omega^2_{Arg,t-1} & \omega_{Arg,t-1}\omega_{Mex,t-1} \\ & \omega^2_{Mex,t-1} \end{bmatrix} B$$

where the most general version assumes no constraints on B (See Engle and Koner [1993]). However, this is usually complicated to estimate and several alternatives have been proposed : First, a popular approach is to restrict B to be diagonal or lower triangular. Second, following Bollerslev [1990], it can be assumed that the conditional moments have a constant correlation. Thus,

$$h_{ArgMex,t} = \rho_{ArgMex}\sqrt{h_{Arg,t}} \cdot \sqrt{h_{Mex,t}},$$

for some fixed parameter ρ_{ArgMex}. Three, as was proposed by Bollerslev, Engle, and Wooldridge [1988], a *vech* formulation could be adopted. Hence, the conditional moments are given by:

$$\begin{bmatrix} h_{Arg,t} \\ h_{ArgMex,t} \\ h_{Mex,t} \end{bmatrix} = \begin{bmatrix} \hat{\zeta}_{Arg} \\ \hat{\zeta}_{ArgMex} \\ \hat{\zeta}_{Mex} \end{bmatrix} + B \begin{bmatrix} \omega^2_{Arg,t-1} \\ \omega_{Arg,t-1}\omega_{Mex,t-1} \\ \omega^2_{Mex,t-1} \end{bmatrix}.$$

Finally, Diebold and Nerlove [1989] and Engle, Ng, and Rothschild [1990] introduced a latent factor ARCH model. that can be estimated

All these techniques impose some type of constraint on the matrix B is order to simplify the estimation procedure. In this paper, the restrictions on B arise from the structural form equations (1) to (6). In particular, the assumption of zero correlation of the structural equation implies that B (in the *vech* specification) has a column of zeros, and that the other elements are given by equation (11).

Finally, it is important to highlight that the covariance restriction on the structural shocks together with the model of heteroskedasticity of the structural residuals impose important constraints on how the reduced form heteroskedasticity can evolve. It is the combination of both aspects what allows for the identification of the structural parameters.[8] On the one hand, in the

[8] There is a new literature studying identification under heteroskedasticity. Most of that literature studies the case

absence of heteroskedasticity, the system would be underidentified. Even if the structural shocks are uncorrelated the structural parameters cannot be recovered from the reduced form estimates. This is indeed the typical problem of simultaneous equations. In fact, in most macro applications the zero correlation on structural residuals is usually imposed, and still the system needs further restrictions. On the other hand, if the covariance of the structural shocks is unrestricted, then the presence of the heteroskedasticity does not add any constraint to the system of equations.

3.2 Argentina and Mexico's sovereign debt

In this section, I estimate the propagation of shocks between Argentina and Mexico in the two sub-samples. The objective is to compare the parameters and understand in which dimensions the transmission of shocks has been modified by the Mexican upgrade.

The model is an extension of the structural ARCH described before, where I have included lags in the mean equation. This does not change the conditions nor the assumptions to achieve identification. The specification is as follows,

$$
\begin{aligned}
Arg_t &= \beta Mex_t + \phi_a X_{t-L} + \varepsilon_t, \\
Mex_t &= \alpha Arg_t + \phi_m X_{t-L} + \eta_t,
\end{aligned}
$$

where X_{t-L} is the vector of L lags yields. In this model the residuals of the reduced form ARCH have the same structure as equations (7) to (11).

The parameters of interest are α, β, and λ. α represents the propagation of shocks, in the mean equation, from Argentina to Mexico, while β is the contemporaneous transmission from Mexico to Argentina. λ are the coefficients describing the diffusion through the second moments. In particular the elements in the diagonal represent the pure country specific ARCH effect, while the elements off the diagonal are the cross country effects.[9]

The data for the non-investment grade period corresponds to the daily EMBI+ yields of Argentina and Mexico from January 1996 to March 6th, 2000. The investment grade sample covers from March 7th, 2000 until May 28th of 2001.[10] I compute the reduced form ARCH using 5 lags

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of unconditional heteroskedasticity. In this paper, I extend those procedure to the more general case of conditional heteroskedasticity.conditional. See Fiorentini and Sentana [1999], Klein and Vella [2000a, b] and Rigobon [2000a].

[9] Additionally, it would be possible to estimate ϕ_a and ϕ_m. However, I am not interested in the dynamic adjustment of the shocks. I am mainly concerned with the estimation of the contemporaneous parameters.

[10] The beginning and end points of the non-investment grade sample were changed for shorter periods, and to exclude March at all. No qualitative difference in the results were found; and therefore, I report only the results using all the sample.

in the mean equation, and restricting the conditional moments by using equations (9) and (11). I estimate the model by GMM.[11]

In Table 3, the results from estimating the non-investment grade sample are shown. The first column is the point estimate, the second column is the standard deviation computed using the asymptotic distribution, and, the third column are the simple z-statistics.

	Point Estimate	St Dev	Z stat
β	0.561583	0.214568	2.6
α	0.425927	0.062756	6.8
$\lambda_{\varepsilon\varepsilon}$	0.976805	0.128062	7.6
$\lambda_{\varepsilon\eta}$	0.002186	0.044164	0.0
$\lambda_{\eta\varepsilon}$	0.013289	0.185224	0.1
$\lambda_{\eta\eta}$	0.969390	0.170360	5.7
ζ_{ε}	0.065374	0.278235	0.2
ζ_{η}	0.072026	0.346842	0.2

Table 3: Estimated Coefficients for the Non-Investment Grade Sample.

The estimates show that 56 percent of the innovations in Mexico sovereign debt are contemporaneously transmitted to Argentina, and that 43 percent of the Argentinean innovations are spread to Mexico. Both estimates are statistically significant and economically relevant. Indeed, these coefficients mean that the minimum possible correlation between Argentina and Mexico is 78.9 percent.

It could be argued that the coefficients are large because in the specification there is an omitted variable that is positively correlated with both country risks. However, remember that in the construction of the data exchange rate and US interest rates risks have been taken into consideration. Therefore, the only unobservable is changes in risk preferences in this particular market that are not reflected in the US interest rate. Still, by any measure, this degree of co-movement is extremely large.

The next four coefficients in Table 3 are the propagation of shocks by the variances. As can be seen, there is a strong country effect (both $\lambda_{\varepsilon\varepsilon}$ and $\lambda_{\eta\eta}$ are close to one), while the cross market coefficients are relatively small: $\lambda_{\varepsilon\eta}$ and $\lambda_{\eta\varepsilon}$ are not statistically different from zero.

In Table 4, the estimates during the investment grade sample are presented. The interpretation is the same as before.

[11] I also implemented an indirect estimation technique as the one described in Duffie and Singelton [1993], Dungey and Martin [2001], Gallant and Tauchen [1996] and Gourieoux, Monfort, and Renault [1993]. The point estimates were very close to the ones obtained here.

	Point Estimate	St Dev	Z stat
β	0.209725	0.092377	2.3
α	0.315230	0.041809	7.5
$\lambda_{\varepsilon\varepsilon}$	0.947990	0.017656	53.7
$\lambda_{\varepsilon\eta}$	0.007266	0.026628	0.3
$\lambda_{\eta\varepsilon}$	0.000000	0.006509	0.0
$\lambda_{\eta\eta}$	0.904422	0.258121	3.5
ζ_{ε}	0.360059	0.068314	5.3
ζ_{η}	0.558291	0.092757	6.0

Table 4: Estimated Coefficients for the Investment Grade Sample.

First, observe that the contemporaneous coefficients are smaller than those estimated during the non-investment grade period. The diffusion of shocks from Mexico to Argentina is only 21 percent, while the transmission from Argentina to Mexico is 32 percent. These coefficients imply that the minimum possible correlation between the countries would be 48.2 percent. Which is substantially smaller than the one found in the previous sample.

Second, the propagation of shocks through the variances have point estimates that are close to those obtained before.

Finally, tests of equality of the two mean coefficients across specifications can be performed. Using the asymptotic distribution, it is possible to reject the hypothesis of joint equality at 95 percent confidence. In conclusion, this evidence suggests that the upgrade to investment grade reduced the propagation of shocks in the mean equation (α and β drop), while the transmission of volatility shocks remained relatively constant.

4 Investor Universe Co-movement

In this section, I analyze the implications of the previous estimated parameters on the co-movement of sovereign bonds. The objective is to study the predicted fall in correlations due to the swing in the investors universe.

The base of comparison is the unconditional correlations. 10000 draws of the structural shocks are generated, and using the estimated coefficients in each of the sub-samples, the new data is constructed and the implied unconditional correlations are computed. The results of this exercise are presented in Table 5.

The first row is the variance of Argentina, the second row is the covariance between Argentina

	NIG	IG
$Var(Arg)$	1.2603	5.3746
$Cov(Arg\,Mex)$	1.0011	2.3252
$Var(Mex)$	1.2358	3.8315
ρ	80%	51%

Table 5: Predicted change in co-movement.

and Mexico, the third row is the variance of Mexican innovations, and the fourth row is the correlation coefficient. All variables are unconditional; meaning that they were computed using all the sample. The first column shows the results from using the non-investment grade coefficients, and the second column correspond the moments obtained by the investment grade estimates.

Several remarks can be extracted from Table 5. First, the predicted unconditional correlations are close to the actual ones shown in Table 1. During the non-investment grade period the unconditional correlation in the data is 88 percent, while the predicted one is 80 percent. In the investment grade sample, the predicted correlation is 51 percent and the sample one is 44. Obviously these are within sample comparisons, but the model fits remarkably well these unconditional correlations.

Second, 36 percent of the correlation is explained by the change in the investors universe. Note that the reduction in unconditional correlations is 29 percent $(81-51)$, which as a percentage of the non-investment grade unconditional correlation is more than a third. Furthermore, this reduction would have been even larger if markets were calm. The fact that Argentina is going through a major crisis implies that the variance of both countries (and their covariance) are larger, biasing upward the correlation estimates.

This evidence indicates that a sizeable proportion of the co-movement across markets is due to the segmentation of the investors universe. It could be argued that the reduction in the correlation is because during the non-investment grade sample several crises increased the co-movement between Argentina and Mexico. However, this is not the case.

These estimates are qualitatively the same as those obtained if only one year of the non-investment grade sample is used. The reason to perform this exercise is that from March 1999 to March 2000 there has not been reported a liquidity shock in the bond market. The estimates of this sub-sample are shown in Table 6.

Note that the coefficients are very close to those in Table 3. In fact, the null hypothesis that the coefficients are the same cannot be rejected at a reasonable confidence (at least 85 percent confidence). These coefficients imply an unconditional correlation of 77 percent, and a minimum possible correlation of 72 percent. Results that are strikingly close to those using the whole sample.

	Point Estimate	St Dev	Z stat
β	0.503166	0.049316	10.2
α	0.358977	0.029395	12.2
$\lambda_{\varepsilon\varepsilon}$	0.973252	0.046860	20.8
$\lambda_{\varepsilon\eta}$	0.000000	0.012284	0.0
$\lambda_{\eta\varepsilon}$	0.000655	0.016057	0.0
$\lambda_{\eta\eta}$	0.866681	0.024735	35.0
ζ_{ε}	0.173041	0.079482	2.2
ζ_{η}	0.555369	0.099131	5.6

Table 6: Estimated Coefficients for the Non-Investment Grade Sample. Using only one year of data.

The implication of these results is that the reduction in the correlation cannot be explained because during the non-investment grade there were more common shocks than afterwards. Indeed, if something has happen is the opposite. During the investment grade sample the emerging economies have been subject to Turkey and Argentina's turmoils. Which reverses the logic.

Hence, the results point to the conclusion that in smaller and segmented investor's markets, wealth shocks are likely to be more relevant in explaining the observed co-movement.

5 Conclusions

This paper examines the properties of the transmission of shocks between Mexico and Argentina before and after Mexico was upgraded from non-investment grade to investment grade. The contributions of the paper are twofold: The first one is related to the relationship between Mexico and Argentina, and the second one is methodological.

I show that the upgrade implied a significant regime shift. The estimates of the contemporaneous relationship between Argentinean and Mexican sovereign bonds indicate that most of the change occurred in the propagation of shocks through the mean equation, while the diffusion through the second moments remained relatively stable. Finally, the last section argues that more than one third of the co-movement has been reduced by the upgrade. The evidence provided in this paper suggests that the change in the investor's universe is the cause of the reduction in co-movement. The paper argues that even though there has not been a liquidity shock or wealth shock to investors that could explain the lack of co-movement (recently), the change in investors universe makes recent wealth shocks less important. Further research should continue exploring this channel.

From the methodological point of view, the paper offers a new identification procedure to solve the problem of simultaneous equations when the data suffers from conditional heteroskedasticity. The present paper is a first pass to the estimation of this type of structural ARCH models. And clearly the model describing the bond yields is very simple. Future research should generalize the present specification to deal with more complex setups.

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A Derivation of the reduced form ARCH.

From the structural form, the reduced form residuals are given by

$$\begin{aligned}\omega_{Arg,t} &= \frac{1}{1-\alpha\beta}\left(\beta\eta_t+\varepsilon_t\right)\\ \omega_{Mex,t} &= \frac{1}{1-\alpha\beta}\left(\eta_t+\alpha\varepsilon_t\right)\end{aligned}$$

where $\omega_{Arg,t}$ and $\omega_{Mex,t}$ are the reduced form innovations. Define the vector of reduced form residuals as $\omega_t \equiv (\omega_{Arg,t}\ \omega_{Mex,t})'$ and the vector of structural shocks as $\sigma_t \equiv (\eta_t\ \varepsilon_t)'$. Denote the conditional covariance matrix as $\Sigma_{\omega,t}$. Where it can be estimated from the sample. Denote η_t^2 and ε_t^2 as the conditional structural moments of the structural shocks. Given equations (1) and (2) the moments of the reduced forms can be written as

$$\begin{aligned}\omega^2_{Arg,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\beta^2\eta_t^2+\varepsilon_t^2\right), &(12)\\ \omega_{Arg,t}\omega_{Mex,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\beta\eta_t^2+\alpha\varepsilon_t^2\right),\\ \omega^2_{Mex,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\eta_t^2+\alpha^2\varepsilon_t^2\right). &(13)\end{aligned}$$

In order to derive the reduced form ARCH, I derive a *vech* specification in the spirit of Bollerslev, Engle, and Wooldridge [1988]. After taking expectations the conditional reduced form residuals can be written in terms of $h_{\varepsilon,t}$, and $h_{\eta,t}$. Define $h_{Arg,t} \equiv E\omega^2_{Arg,t}$, $h_{ArgMex,t} \equiv E\omega_{Arg,t}\omega_{Mex,t}$, and $h_{Mex,t} \equiv E\omega^2_{Mex,t}$, then

$$\begin{aligned}h_{Arg,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\beta^2 h_{\eta,t}+h_{\varepsilon,t}\right),\\ h_{ArgMex,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\beta h_{\eta,t}+\alpha h_{\varepsilon,t}\right),\\ h_{Mex,t} &= \frac{1}{(1-\alpha\beta)^2}\left(h_{\eta,t}+\alpha^2 h_{\varepsilon,t}\right).\end{aligned}$$

Using equations (5) and (6) the expected conditional moments of the structural shocks can be written in terms of the structural shock realizations. Therefore, the conditional moments are

$$\begin{aligned}h_{Arg,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\zeta_{Arg}+\varepsilon^2_{t-1}\left[\beta^2\lambda_{\eta\varepsilon}+\lambda_{\varepsilon\varepsilon}\right]+\eta^2_{t-1}\left[\beta^2\lambda_{\eta\eta}+\lambda_{\varepsilon\eta}\right]\right),\\ h_{ArgMex,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\zeta_{MexArg}+\varepsilon^2_{t-1}\left[\beta\lambda_{\eta\varepsilon}+\alpha\lambda_{\varepsilon\varepsilon}\right]+\eta^2_{t-1}\left[\beta\lambda_{\eta\eta}+\alpha\lambda_{\varepsilon\eta}\right]\right),\\ h_{Mex,t} &= \frac{1}{(1-\alpha\beta)^2}\left(\zeta_{Mex}+\varepsilon^2_{t-1}\left[\lambda_{\eta\varepsilon}+\alpha^2\lambda_{\varepsilon\varepsilon}\right]+\eta^2_{t-1}\left[\lambda_{\eta\eta}+\alpha^2\lambda_{\varepsilon\eta}\right]\right).\end{aligned}$$

Finally, ε^2_{t-1} and η^2_{t-1} can be written as a function of only two of the moments of the reduced form residuals: $\omega^2_{Arg,t-1}$, $\omega^2_{Mex,t-1}$ and $\omega_{Arg,t-1}\omega_{Mex,t-1}$. Using the variances of the reduced form residuals,

$$\begin{aligned}
\varepsilon^2_{t-1} &= \frac{1-\alpha\beta}{1+\alpha\beta}\left[\omega^2_{Arg,t-1}-\beta^2\omega^2_{Mex,t-1}\right], \\
\eta^2_{t-1} &= \frac{1-\alpha\beta}{1+\alpha\beta}\left[-\alpha^2\omega^2_{Arg,t-1}+\omega^2_{Mex,t-1}\right].
\end{aligned}$$

To write the structural conditional moments with only two of the reduced form conditional moments can be performed because the zero covariance restriction has been imposed. Conversely, if the covariance is unrestricted, then the moments of the structural form residuals are three, and require all three reduced form moments to been able to write them.

In conclusion, the ARCH structure is as follows:

$$\begin{bmatrix} h_{Arg,t} \\ h_{ArgMex,t} \\ h_{Mex,t} \end{bmatrix} = \begin{bmatrix} \hat{\zeta}_{Arg} \\ \hat{\zeta}_{ArgMex} \\ \hat{\zeta}_{Mex} \end{bmatrix} + \frac{1}{1-(\alpha\beta)^2} A \begin{bmatrix} \omega^2_{Arg,t-1} \\ \omega^2_{Mex,t-1} \end{bmatrix},$$

where A is a three by two matrix given by

$$A \equiv \begin{bmatrix} \left[\beta^2\lambda_{\eta\varepsilon}+\lambda_{\varepsilon\varepsilon}\right]-\alpha^2\left[\beta^2\lambda_{\eta\eta}+\lambda_{\varepsilon\eta}\right] & -\beta^2\left[\beta^2\lambda_{\eta\varepsilon}+\lambda_{\varepsilon\varepsilon}\right]+\left[\beta^2\lambda_{\eta\eta}+\lambda_{\varepsilon\eta}\right] \\ \left[\beta\lambda_{\eta\varepsilon}+\alpha\lambda_{\varepsilon\varepsilon}\right]-\alpha^2\left[\beta\lambda_{\eta\eta}+\alpha\lambda_{\varepsilon\eta}\right] & -\beta^2\left[\beta\lambda_{\eta\varepsilon}+\alpha\lambda_{\varepsilon\varepsilon}\right]+\left[\beta\lambda_{\eta\eta}+\alpha\lambda_{\varepsilon\eta}\right] \\ \left[\lambda_{\eta\varepsilon}+\alpha^2\lambda_{\varepsilon\varepsilon}\right]-\alpha^2\left[\lambda_{\eta\eta}+\alpha^2\lambda_{\varepsilon\eta}\right] & -\beta^2\left[\lambda_{\eta\varepsilon}+\alpha^2\lambda_{\varepsilon\varepsilon}\right]+\left[\lambda_{\eta\eta}+\alpha^2\lambda_{\varepsilon\eta}\right] \end{bmatrix}.$$

and where the constants are given by

$$\begin{aligned}
\hat{\zeta}_{Arg} &= \frac{1}{(1-\alpha\beta)^2}\left(\beta^2\zeta_\eta+\zeta_\varepsilon\right), \\
\hat{\zeta}_{ArgMex} &= \frac{1}{(1-\alpha\beta)^2}\left(\beta\zeta_\eta+\alpha\zeta_\varepsilon\right), \\
\hat{\zeta}_{Mex} &= \frac{1}{(1-\alpha\beta)^2}\left(\zeta_\eta+\alpha^2\zeta_\varepsilon\right).
\end{aligned}$$

B Lack of identification under an unconstrained structural model.

This section studies the case in which a common heteroskedastic shock is included in the specification. It shows that identification of the structural parameters is not achieved.

Assume that the structural model is:

$$\begin{aligned} Arg_t &= \beta Mex_t + \gamma z_t + \varepsilon_t, \\ Mex_t &= \alpha Arg_t + z_t + \eta_t, \end{aligned}$$

where z_t is a common shock. Assume the structural shocks satisfy the following ARCH model:

$$\begin{aligned} \varepsilon_t &= \sqrt{h_{\varepsilon,t}} \cdot v_{\varepsilon,t} \\ \eta_t &= \sqrt{h_{\eta,t}} \cdot v_{\eta,t} \\ z_t &= \sqrt{h_{z,t}} \cdot v_{z,t} \end{aligned}$$

where $v_{\varepsilon,t}$, $v_{\eta,t}$, and $v_{z,t}$ are uncorrelated shocks, with mean zero and variance one. The structural ARCH is

$$\begin{pmatrix} h_{\varepsilon,t} \\ h_{\eta,t} \\ h_{z,t} \end{pmatrix} = \begin{pmatrix} \zeta_\varepsilon \\ \zeta_\eta \\ \zeta_z \end{pmatrix} + \lambda \begin{pmatrix} \varepsilon_{t-1}^2 \\ \eta_{t-1}^2 \\ z_{t-1}^2 \end{pmatrix},$$

where ζ_ε, ζ_η, and ζ_z are positive constants and

$$\lambda \equiv \begin{bmatrix} \lambda_{\varepsilon\varepsilon} & \lambda_{\varepsilon\eta} & \lambda_{\varepsilon z} \\ \lambda_{\eta\varepsilon} & \lambda_{\eta\eta} & \lambda_{\eta z} \\ \lambda_{z\varepsilon} & \lambda_{z\eta} & \lambda_{zz} \end{bmatrix}.$$

The inclusion of a heteroskedastic common shock is equivalent to assume no restrictions on the covariance between ε_t and η_t. This model is simpler from the expositional point of view.

The reduced form is the following

$$\begin{aligned} \omega_{Arg,t} &= \frac{1}{1-\alpha\beta} \left(\beta\eta_t + \varepsilon_t + (\beta+\gamma) z_t \right) \\ \omega_{Mex,t} &= \frac{1}{1-\alpha\beta} \left(\eta_t + \alpha\varepsilon_t + (1+\alpha\gamma) z_t \right) \end{aligned}$$

where $\omega_{Arg,t}$ and $\omega_{Mex,t}$ are the reduced form innovations. From the structural equations, the expected conditional moments of the reduced form residuals are:

$$\begin{aligned} h_{Arg,t} &= \frac{1}{(1-\alpha\beta)^2} \left(\beta^2\eta_t^2 + \varepsilon_t^2 + (\beta+\gamma)^2 z_t^2 \right), \\ h_{ArgMex,t} &= \frac{1}{(1-\alpha\beta)^2} \left(\beta\eta_t^2 + \alpha\varepsilon_t^2 + (1+\alpha\gamma)(\beta+\gamma) z_t^2 \right), \\ h_{Mex,t} &= \frac{1}{(1-\alpha\beta)^2} \left(\eta_t^2 + \alpha^2\varepsilon_t^2 + (1+\alpha\gamma)^2 z_t^2 \right), \end{aligned}$$

which after taking expectations can be written in terms of $h_{\varepsilon,t}$, $h_{\eta,t}$, and $h_{z,t}$ as follows,

$$\begin{pmatrix} h_{Arg,t} \\ h_{ArgMex,t} \\ h_{Mex,t} \end{pmatrix} = \frac{1}{(1-\alpha\beta)^2} \begin{pmatrix} 1 & \beta^2 & (\beta+\gamma)^2 \\ \alpha & \beta & (1+\alpha\gamma)(\beta+\gamma) \\ \alpha^2 & 1 & (1+\alpha\gamma)^2 \end{pmatrix} \begin{pmatrix} h_{\varepsilon,t} \\ h_{\eta,t} \\ h_{z,t} \end{pmatrix}$$

Which using the ARCH equations the conditional moments are

$$\begin{pmatrix} h_{Arg,t} \\ h_{ArgMex,t} \\ h_{Mex,t} \end{pmatrix} = \begin{pmatrix} \hat{\zeta}_\varepsilon \\ \hat{\zeta}_\eta \\ \hat{\zeta}_z \end{pmatrix} + \frac{1}{(1-\alpha\beta)^2} \begin{pmatrix} 1 & \beta^2 & (\beta+\gamma)^2 \\ \alpha & \beta & (1+\alpha\gamma)(\beta+\gamma) \\ \alpha^2 & 1 & (1+\alpha\gamma)^2 \end{pmatrix} \lambda \begin{pmatrix} \varepsilon_{t-1}^2 \\ \eta_{t-1}^2 \\ z_{t-1}^2 \end{pmatrix}$$

Finally, using the reduced form equations the structural errors ε_{t-1}^2, η_{t-1}^2, and z_{t-1}^2 can be written in terms of $\omega_{Arg,t-1}^2$, $\omega_{Arg,t-1}\omega_{Mex,t-1}$, and $\omega_{Mex,t-1}^2$. This implies that the reduced form ARCH model can be written as

$$\begin{pmatrix} h_{Arg,t} \\ h_{ArgMex,t} \\ h_{Mex,t} \end{pmatrix} = C + B \begin{pmatrix} \omega_{Arg,t-1}^2 \\ \omega_{Arg,t-1}\omega_{Mex,t-1} \\ \omega_{Mex,t-1}^2 \end{pmatrix}.$$

where B is a three by three matrix.

Note that from the reduced form the number of estimable parameters from B is nine, while the number of structural form coefficients is 12: three from the mean equations (α, β, γ) plus nine from the ARCH model (λ). Thus, the system is underidentified.

As was indicated above, the reason why the identification is not achieved in this case is because the structural assumptions impose no constraints on the reduced form. Note that indeed, the reduced form obtained here is the unrestricted *vech* setup

C Equivalence between the supply-demand specification and the latent factor model.

The structural form assumed is the following:

$$
\begin{aligned}
Arg_t &= \beta Mex_t + \varepsilon_t, \\
Mex_t &= \alpha Arg_t + \eta_t,
\end{aligned}
$$

several papers have adopted a latent factor model as follows:

$$
\begin{aligned}
Arg_t &= \gamma_a \eta_t + \varepsilon_t, \\
Mex_t &= \gamma_m \varepsilon_t + \eta_t.
\end{aligned}
$$

The two models are equivalent. Solving for the country yields in the first, the following reduced form is found:

$$
\begin{aligned}
Arg_t &= \frac{1}{1-\alpha\beta}\left(\beta\eta_t + \varepsilon_t\right), \\
Mex_t &= \frac{1}{1-\alpha\beta}\left(\eta_t + \alpha\varepsilon_t\right).
\end{aligned}
$$

Note that if the variance of the structural shocks are normalized by $\frac{1}{1-\alpha\beta}$, then $\gamma_a = \beta$, and $\gamma_m = \alpha$. Therefore, the two models produce the exact same reduced form implications, and even the interpretation of the coefficients is the same. The only difference is in the size of the variance of the identified shocks.

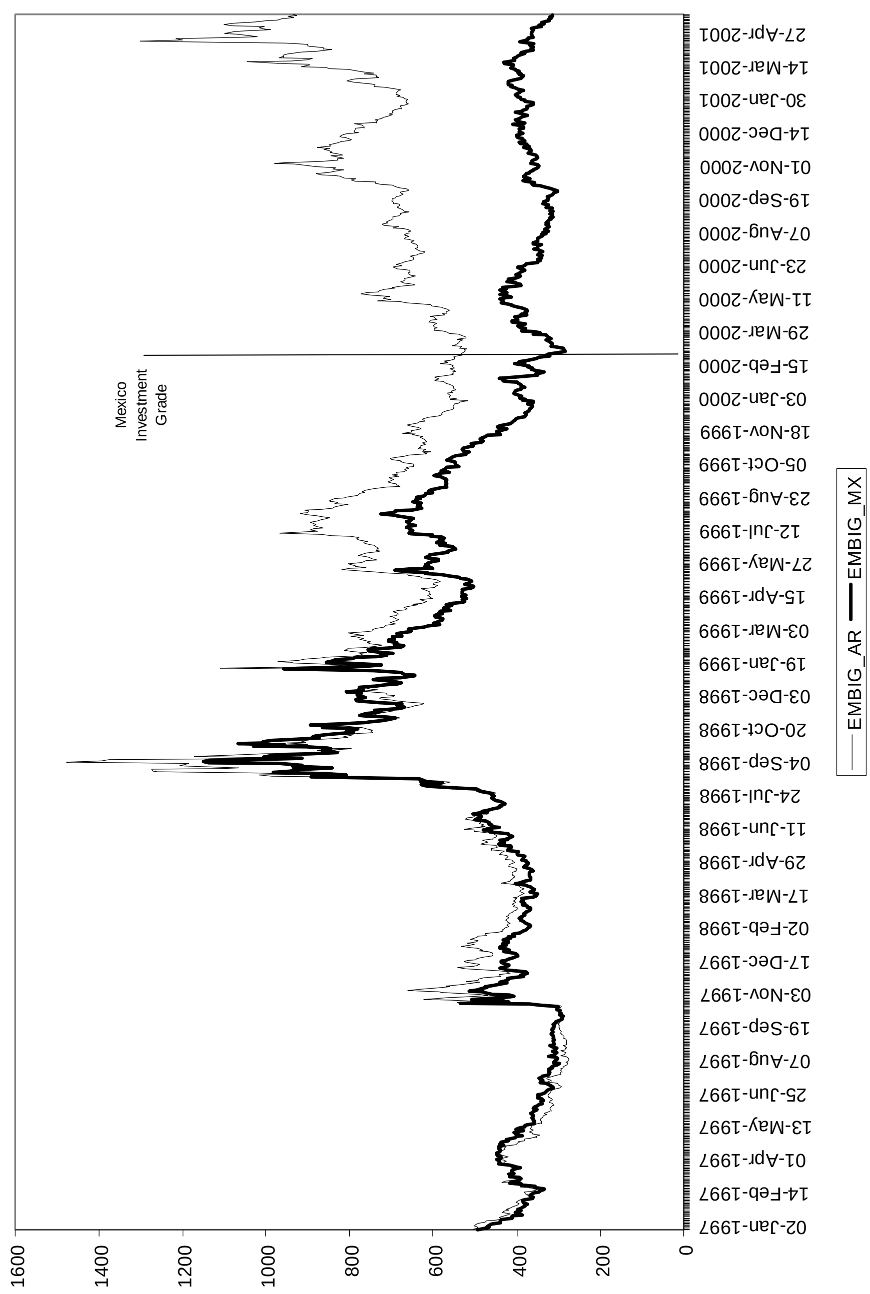

Figure 1: Brady Bond Yields

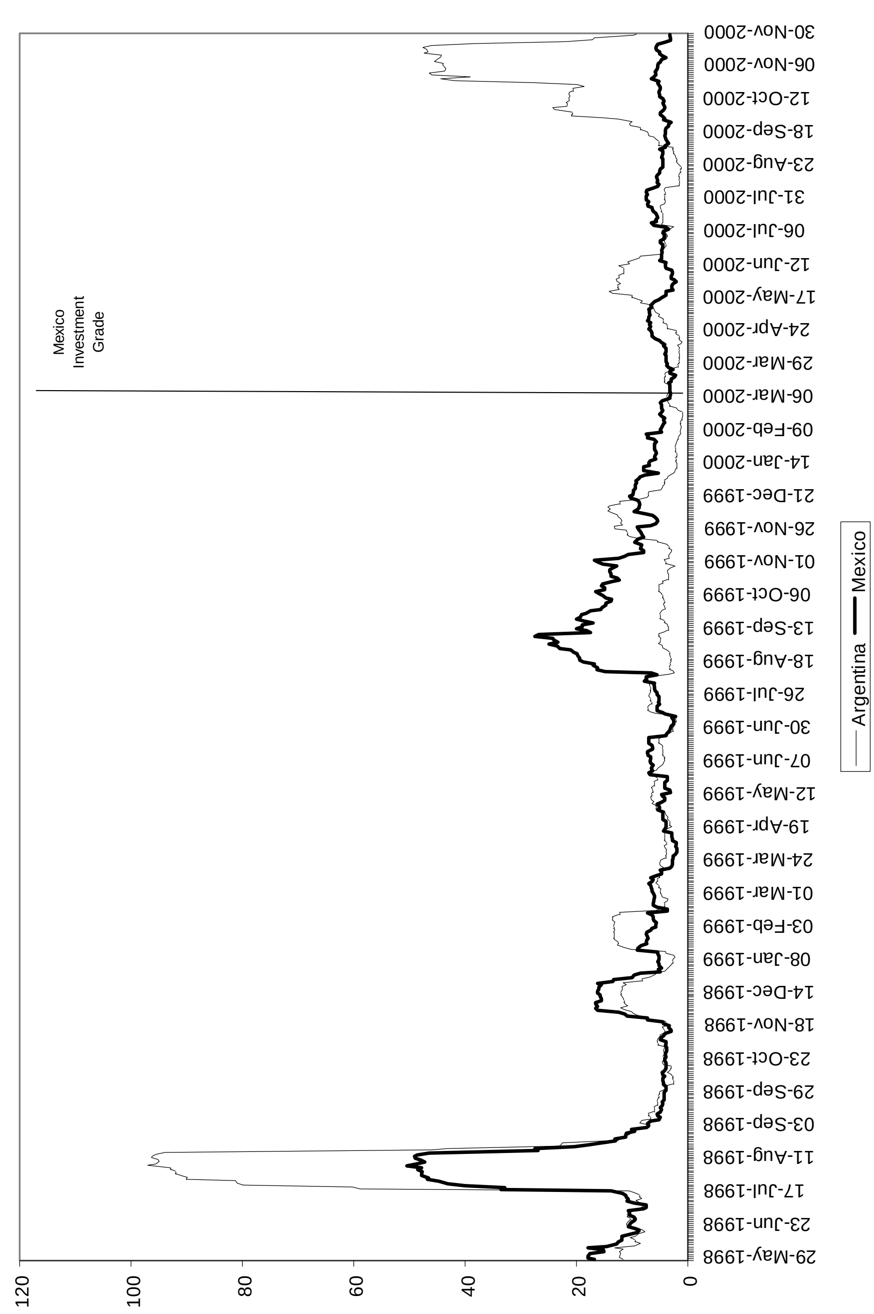

Figure 2: Rolling Variance of Mexican and Argentinean Yields.

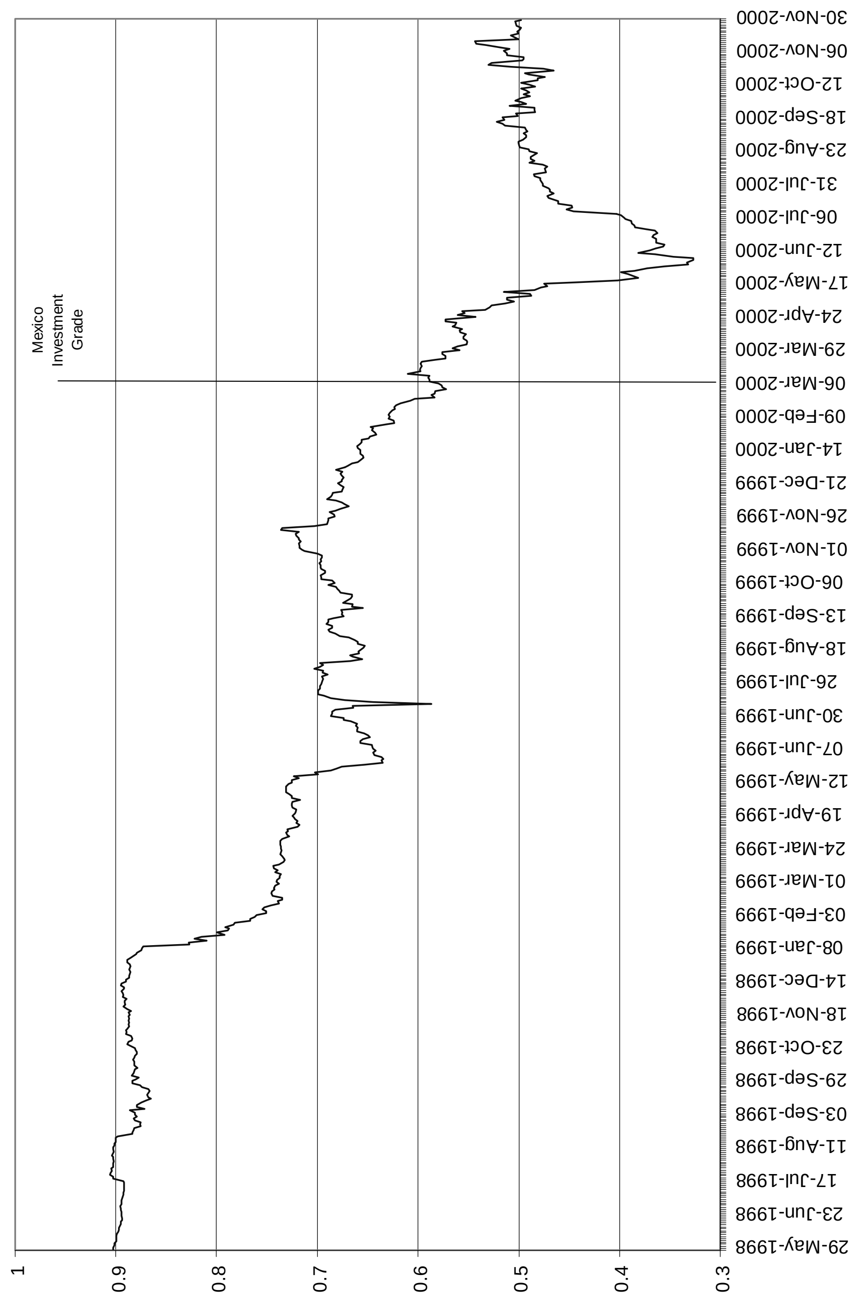

Figure 3: Rolling correlation between Argentinean and Mexican yields. Rolling window of 90 days.