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Sovereign Bond Prices, Haircuts and Maturity

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Tamon Asonuma, Dirk Niepelt, and Romain Rancière
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ABSTRACT

Rejecting a common assumption in the sovereign debt literature, we document that creditor losses ("haircuts") during sovereign restructuring episodes are asymmetric across debt instruments. We code a comprehensive dataset on instrument-specific haircuts for 28 debt restructurings with private creditors in 1999-2015 and find that haircuts on shorter-term debt are larger than those on debt of longer maturity. In a standard asset pricing model, we show that increasing short-run default risk in the run-up to a restructuring episode can explain the stylized fact. The data confirms the predicted relation between perceived default risk, bond prices, and haircuts by maturity.

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1 Introduction

Conventional wisdom among practitioners and academics holds that the treatment of private creditors in sovereign debt exchanges does not vary by the maturity of instruments they hold (Institute of International Finance 2012; 2015).[1] After all, the pari passu clause, which is commonly included in unsecured cross-border corporate and international sovereign debt contracts, provides that the debt instruments issued under such contracts will rank equally among themselves and with all other present or future unsubordinated and unsecured external debt obligations of the borrower (IMF 2014). In line with this view, the empirical literature on sovereign debt restructurings has documented that creditor losses ("haircuts") tend to be symmetric across private debt instruments—with a few exceptional restructuring episodes (Sturzenegger and Zettelmeyer 2006, 2008).

In this paper, we challenge the conventional wisdom. Based on a new, comprehensive dataset on instrument-specific haircuts in 28 sovereign debt restructuring episodes (18 external and 10 domestic) with private creditors over the period 1999–2015, we document that creditor losses systematically vary by maturity of debt instrument. Specifically, we show that haircuts on shorter-term debt tend to be larger.

Our finding is robust. It holds when we measure haircuts according to the standard Sturzenegger and Zettelmeyer (2006, 2008) measure which is computed based both on a market price and a hypothetical net present value. It also holds when we measure haircuts alternatively, based on the price of the "old" instrument before the exchange relative to the price of the "new" instrument immediately after the restructuring. Our alternative measure can easily be computed because it only requires observed market prices. Moreover, it can be computed for windows of different length before the actual restructuring date.

Our finding also is robust when we control for characteristics of the exchanged bonds or the restructuring episode. For instance, it holds both in external and domestic restructuring episodes; it holds when restructurings occur preemptively or post default[2]; and it holds independently of the exchange method that is, whether the exchange involves a single menu of new instruments or different menus depending on type of restructured bonds.[3] Cross-sectional regressions controlling for instrument- and episode-specific effects robustly confirm the stylized fact.

Digging deeper, we find that the negative relation between the maturity of a bond and the haircut suffered during a debt exchange reflects the fact that ceteris paribus, short-term bond prices tend to be higher before an exchange than the prices of longer-term debt. In other words, in the run-up to a restructuring episode, short-term bond prices tend to converge to long-term

[1] IMF (2015) documents that the terms of exchange often differ along other dimensions including governing law, currency denomination, type of instrument, and residency of debt holder.

[2] Asonuma and Trebesch (2016) define a restructuring as preemptive when the exchange takes place prior to or without a missed payment, and as post-default otherwise.

[3] In the restructuring cases where the exchange involves different menus depending on type of restructured bonds, short-term bond holders are often offered menus which include compensations for estimated creditor losses, i.e. cash payments ("sweeteners").

bond prices from above.

Building on a standard model of arbitrage-free risk-neutral pricing we motivate the stylized fact as the natural implication of the term structure of default risk and its time variation. As long as investors anticipate strictly positive default risk after the maturity date of short-term debt, the price of the latter exceeds the price of longer-term debt (holding all other characteristics constant). Prices of short-term bonds therefore exceed prices of longer-term bonds, and when a restructuring episode ends with all outstanding instruments being exchanged against the same new instrument (or basket of new instruments) then this implies our main stylized fact: A higher haircut on short-term bonds, measured either way. Moreover, if in the run-up to a debt restructuring episode investors perceive increasingly short-term default risk then the prices of short- and longer-term bonds converge. This implies that the extent to which the possible haircuts vary by maturity weakens as the restructuring event approaches. Intuitively, successive bad news in the run-up to the restructuring compress price differentials across maturities and as a consequence, diminish the ex-post haircut differentials as well.

We then explore empirically the theoretical prediction that prices of short- and longer-term bonds converge because of increasingly short-term default risk. In panel regressions covering the period from 6 months before the announcement of a restructuring until the date of exchange or default, and controlling for restructuring-specific factors we find that the data confirms the prediction.

Our paper relates to the empirical literature on creditor losses due to restructurings, specifically Eichengreen and Portes (1986, 1989), Lindert and Morton (1989), Sturzenegger and Zettelmeyer (2006, 2008), Benjamin and Wright (2009), Cruces and Trebesch (2013), Finger and Mecagni (2007), Bedford et al. (2005), as well as Diaz-Cassou et al. (2008).[4] Almost all studies report the average haircuts for the restructuring.[5] Only a few (Sturzenegger and Zettelmeyer 2006, 2008; and Zettelmeyer et al. 2013) document instrument-specific haircuts, but only for selected restructuring episodes. In contrast, we report instrument-specific haircuts for a large sample of restructuring episodes; we document a novel stylized fact concerning the systematic relationship between haircut and maturity; and we offer an explanation for this fact.

Our paper also contributes to the theoretical literature building on Eaton and Gersovitz' (1981) classic framework and studying the maturity structure of sovereign debt for instance Rodrik and Velasco (1999), Jeanne (2009), Broner et al. (2013), Arellano and Ramanarayanan (2012), Chatterjee and Eyigungor (2012), Fernández and Martin (2014), Aguiar and Amador (2013), Hatchondo and Martinez (2009), Hatchondo et al. (2016), and Niepelt (2014). These papers aim at rationalizing why sovereigns choose a specific portfolio composition, often assuming that haircuts after default are symmetric across bonds with different maturities. We document that this assumption is not supported by the data, and we offer an explanation for this new stylized fact which, as we also show, is consistent with other evidence.

The remainder of the paper is structured as follows. Section 2 explains two measures of

[4] Cline (1995) and Rieffel (2003) contain further haircut estimates for several episodes.

[5] Applying a similar approach, Hatchondo et al. (2014) measure average capital gains on restructured debt.

creditor losses (haircuts) and our dataset. Section 3 documents the stylized facts on haircuts and bond price dynamics. Section 4 presents a theoretical model of sovereign bond prices. Section 5 derives the estimated default probability term structure. A brief conclusion summarizes our discussion.

2 Haircuts

We construct a new dataset covering 28 sovereign debt restructuring episodes over the period 1999–2015. The dataset documents the size of private creditor losses (haircuts) for individual debt instruments based on two measures. First, the well-known Sturzenegger and Zettelmeyer (2006, 2008) measure. And second, the market price of the new debt instrument received in exchange for an old debt security, relative to the price of the old instrument at some date prior to the exchange.

2.1 Haircut Measures

Sturzenegger and Zettelmeyer (2006, 2008, SZ from now on) propose a haircut measure that is widely used among academics and increasingly so among practitioners as well.[6] SZ define the recovery rate (one minus the haircut) as the ratio of the net present values of two cash flow streams, a "new" one and an "old" one. The new cash flow stream reflects the contractually defined payments of the new debt instrument(s) received in the course of a debt exchange; the cash flows are discounted at the yield to maturity of the new instrument(s). The old cash flow stream reflects the contractually defined payments of the old debt instrument that was exchanged; these cash flows are also discounted at the yield to maturity of the new instrument(s).

Let $NPV_t(i, r)$ denote the net present value as of time t of the contractually defined cash flow stream of a debt instrument i when the stream is discounted at the interest rate r. Moreover, let e denote the new, exchanged debt instrument. Formally, the SZ recovery rate measure for instrument i exchanged at date T is defined as

$$SZR_T^i \equiv \frac{NPV_T(e, r_T^e)}{NPV_T(i, r_T^e)} \tag{1}$$

where r_T^e denotes the yield to maturity of instrument e at the time of the exchange (exit yield).

By definition, the price of a bond equals the present discounted value of its contractually defined cash flow stream when the cash flows are discounted at the yield to maturity. Formally, letting p_t^i denote the ex-coupon price of instrument i at date t,

$$p_t^i \equiv NPV_t(i, r_t^i)$$

[6] SZ haircuts differ from "market haircuts" used by financial market participants which compare the market value of new debt and cash received to the sum of *outstanding face value* of the old debt and past due interests (Sturzenegger and Zettelmeyer 2006, 2008).

This implies that the numerator in the SZ recovery rate measure can directly be observed: it is the price of the new, exchanged debt instrument immediately after the exchange. In contrast, the denominator of the measure cannot directly be observed and does not have a meaningful price counterpart: it represents the present discounted value (at the time of the exchange) of the old debt instrument's cash-flow stream when discounted at the new debt instrument's exit yield. Computing the denominator and thus, the SZ recovery rate requires information about the cash flow stream of the old instrument as well as the exit yield of the new instrument.

This information requirement poses a problem if one wishes to assess haircuts for many instruments, as we do. An obvious solution is to replace the net present value in the denominator of the SZ measure by a different net present value that equals an observed market price. This is easily achieved by replacing the exit yield in the net present value formula by the yield to maturity of the old instrument before the exchange. An additional advantage of this approach is that it opens the way for computing haircut measures that are defined over longer "windows."

We therefore define the alternative recovery rate measure, "Exchange recovery rate"

$$R^i_{t,T} \equiv \frac{NPV_T(e, r^e_T)}{NPV_t(i, r^i_t)} \tag{2}$$

which is, by definition, identical to the ratio of two prices. The numerator on the right-hand side of equation (2), which is the same as in equation (1), equals the price of the new instrument immediately after the exchange. The denominator equals the price of the old instrument at some date prior to the exchange, $t \leq T$.

Note that our new measure differs from the SZ measure even for $t = T$. This difference reflects the fact that the two measures use different base values to assess the recovery rate. While SZ use a synthetic value as base value, our measure (computed for $t = T$) uses the market price of the old instrument. In our empirical analysis, we find that quantitatively this difference is minor,

$$\frac{R^i_{T,T}}{SZR^i_T} = \frac{NPV_T(i, r^e_T)}{NPV_T(i, r^i_t)} \approx 1$$

While for a single instrument, the dependence of $R^i_{t,T}$ on window length $T - t$ may be of limited interest, the differential dependence for different instruments does provide useful information. In particular, for two instruments i and j and varying window lengths, $T - t$, the relative recovery rates, $R^i_{t,T} - R^j_{t,T}$, illustrate which instrument's price converged more quickly to the market value of the new instruments received in the exchange.

2.2 Data

Our empirical analysis covers 28 episodes with 18 external and 10 domestic sovereign debt restructurings that involved private creditors, over the period 1999–2015. To date an episode, we rely on Asonuma and Papaioannou (2016) and Asonuma and Trebesch (2016) for domestic and external restructuring episodes, respectively. Following these authors, we define the start

of an episode as the month in which a default occurs or a distressed restructuring is announced; and the end of an episode as the month of the final agreement or of the implementation of the debt exchange.[7]

We construct a novel comprehensive dataset with information about instrument specific haircuts according to the two measures defined previously. For each instrument i, we collect information on its maturity, coupon, payment type, and options. Moreover, we collect the same information for each new instrument e that investors received in exchange for i. We collect this information from several sources: Offering memoranda, press releases from governments, financial sector databases and reports, IMF staff reports, SZ, Cruces and Trebesch (2013), and Asonuma and Papaioannou (2016). For information on market yields of the instruments and on market prices, we rely on financial sector databases such as Bloomberg, Datastream, Dealogic, and JP Morgan Markit.

Table 1 provides information about the scope of our dataset. With data on haircuts according to the SZ measure for 449 instruments in 18 external and 10 domestic sovereign debt restructuring episodes, it contains more than twice the number of observations than SZ. In addition, it contains information on haircuts according to our new measure (Exchange haircuts) for 111 instruments in 12 external and 6 domestic debt restructurings. The smaller number of observations for the new measure is due to the more limited information on bond yields and prices. Where price data is available, the dataset collects information for $R^i_{t,T}$ from 6–9 months prior to the announcement of the restructuring to the exchange in monthly frequency. Appendix A provides an overview over the restructuring episodes included in our sample.

[7]Following Standard & Poor's definition, distressed restructurings involve terms that are less favorable than the original terms of the bonds or loans.

Table 1: Scope of Dataset

	Our data	Sturzenegger and Zettelmeyer (2006, 2008)
SZ Haircuts (Recovery Rates)	**28 restructurings** -18 external -10 domestic **449 instruments** - only at exchange	**11 restructurings** - 8 external - 3 domestic **199 instruments** - only at exchange
Exchange Haircuts (Recovery Rates)	**18 restructurings** -12 external - 6 domestic **119 instruments** - from 6–9 months before the announcement of restructuring to the exchange	

3 New Stylized Facts

Our comprehensive instrument-specific dataset reveals two new stylized facts which are the main findings of the paper. First, using either of the measures defined in Section 2, we document the following: Haircuts on shorter-term bonds tend to be larger than those on bonds of longer maturity. Second, we find that the negative relation between the maturity of a bond and the haircut suffered during a debt exchange reflects the fact that ceteris paribus, short-term bond prices tend to be higher before an exchange than the prices of longer-term debt.

3.1 Haircuts

Stylized fact 1: Haircuts on short-term bonds are larger than those on longer-term bonds.

Figure 1 reports recovery rates according to SZ, SZR_T^i, by maturity for two restructuring episodes: Russia 1998–2000 (external) and Dominica 2003–2004 (domestic). Figure A1 in Appendix B reports two additional episodes: Belize 2006–2007 (external) and Saint Kitts and Nevis 2011–2012 (external). In all these cases and in many more not reported in the Figures, recovery rates on short-term bonds are substantially smaller than those on longer-term bonds. This holds true independently of the type of debt (e.g., external debt in Russia 1998–2000 or Greece 2011–2012 vs. domestic debt in Uruguay 2003 or Cyprus 2013), the restructuring strategy (e.g., preemptive restructuring in Greece 2011–2012 or Cyprus 2013 vs. post-default restructuring in Ecuador 1999–2000 or Russia 1998–2000), or the exchange method (e.g., exchange against a

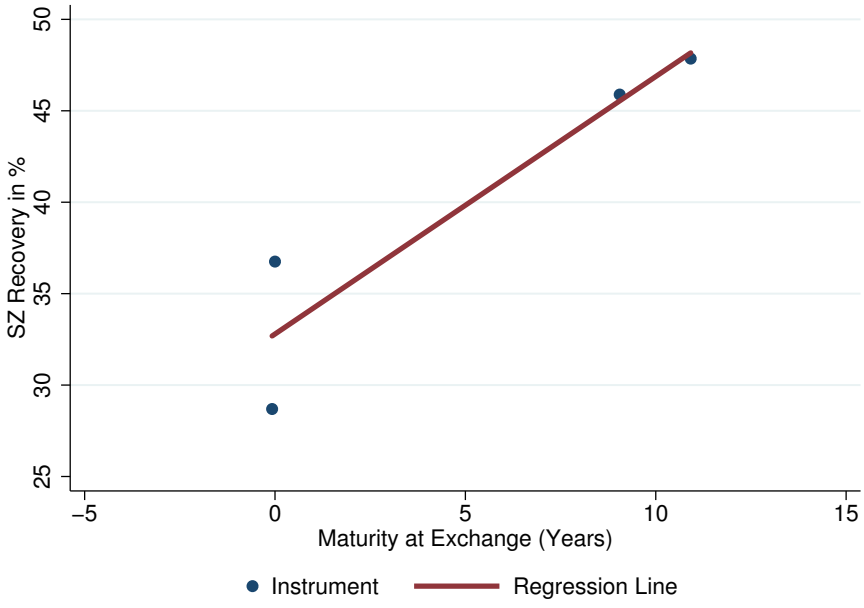
new instrument from a single menu in Greece 2011–2012 or Saint Kitts and Nevis 2011–2012 vs. new instruments from different menus in Cyprus 2013 or Ukraine 2000). In Appendix C, we document that these results are robust to employing modified SZ recovery rate measures based on higher or lower yields (exit yield +/-1 percent).

We obtain the same qualitative finding when we use our alternative recovery rate measure, $R^i_{t,T}$. Figure 2 reports $R^i_{t,T}$ by maturity for two restructuring episodes: Pakistan 1999 (external), and Cyprus 2013 (domestic). Figure A2 in Appendix B reports two additional cases: Greece 2011–2012 (external) and Uruguay 2003 (domestic).[8] Moreover, Figure A3 and Table A2 in Appendix B show that our results remain robust when we evaluate our alternative recovery rate measures at different points in time, i.e. 6 months before and after the announcement of restructurings.

[8] We see the same pattern in the external debt restructuring in Uruguay 2003.

Figure 1: SZ Recovery and Maturity

(i) Russia 1998–2000 External



(ii) Dominican 2003–2004 Domestic

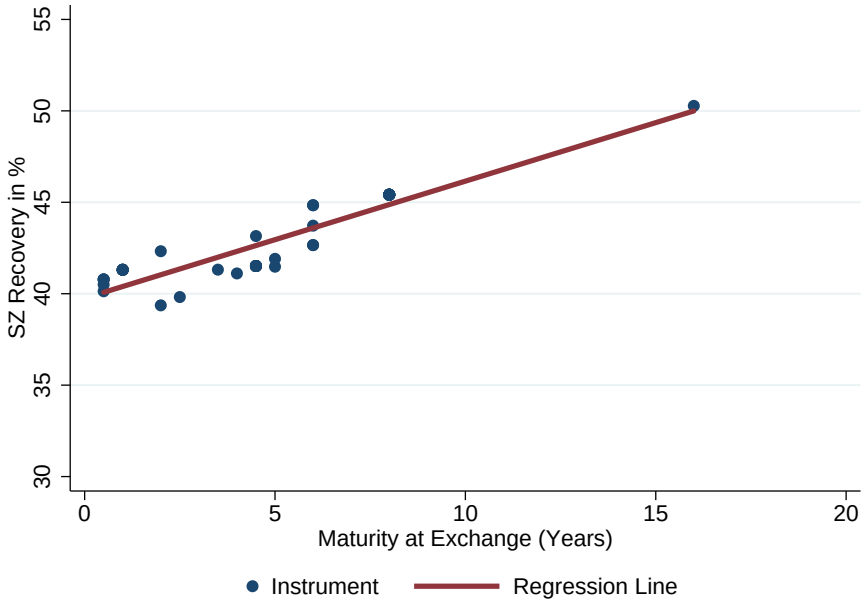
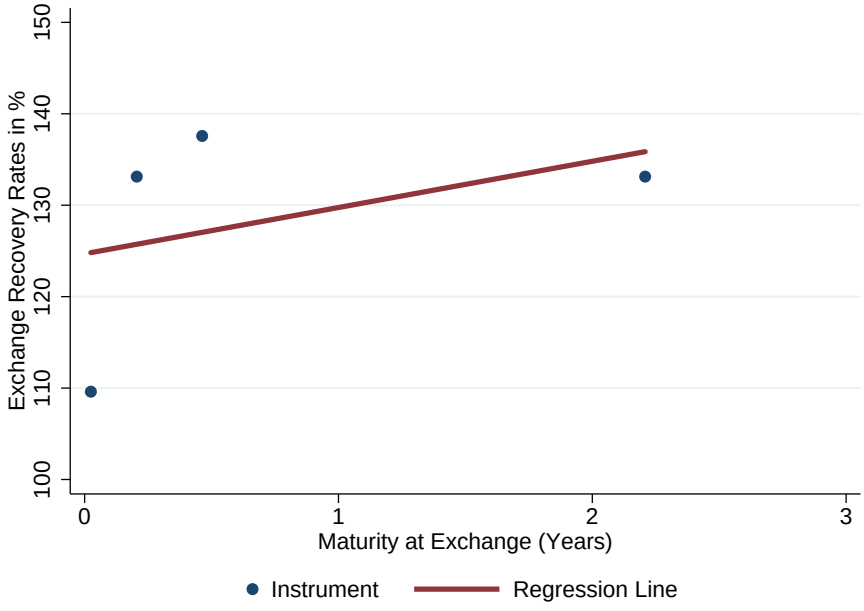


Figure 2: Exchange Recovery and Maturity

(i) Pakistan 1999 External



(ii) Cyprus 2013 Domestic

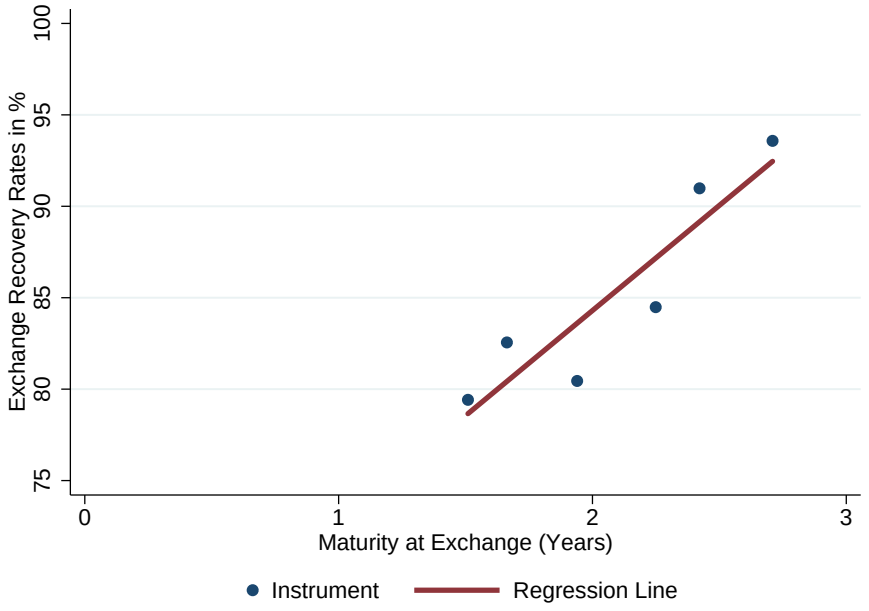


Table 2 provides econometric support for the stylized fact. It reports the results of a cross-sectional regression of the recovery rate (measured according to SZ or our measure) on maturity as well as additional instrument- and episode-specific controls:

$$RecoverRate_j^i = c + \beta_1 Maturity_j^i + \beta_2 x_j^i + \beta_3 y_j + c_j + \epsilon_{i,j} \tag{3}$$

Here, $Maturity_j^i$ denotes the remaining maturity of instrument i at the date of exchange in restructuring episode j; x_j^i is a vector of instrument-specific variables, such as the coupon rate, a dummy in case of a floating coupon rate, a dummy for the type of amortization profile (amortization only at maturity, ["bullet"] vs. already before maturity); y_j is a vector of episode-specific variables, such as the duration of the restructuring and the type of debt (domestic or external); and c_j is an episode-specific fixed effect.

The main result reported in Table 2 is that the effect of maturity on recovery rates (both according to SZ and our measure) is positive and significant at the 1-percent level. Quantitatively, the effect is significant as well: on average, the recovery rate on a 10-year bond is 3–12 percentage points higher than on a 1-year bond. Columns (2) and (4) also reveal the effects of other controls. Instruments with high fixed coupon rates or with floating rates experience smaller recovery rates. The sign of the effect of the amortization profile is not precisely estimated. Because of different sample sizes, the effect of restructuring duration on the recovery rate only is identified in specification (2); the former is positive and significant.

Table 2: Cross-sectional Regression Results

	SZ Recovery Rate	SZ Recovery Rate, with controls	**Exchange Recovery Rate (announcement)**	Exchange Recovery Rate, with controls
	(1)	(2)	(3)	(4)
	coef/se	coef/se	coef/se	coef/se
Maturity of Instrument (years)	0.30***	0.35***	1.03***	1.24***
	(0.08)	(0.08)	(0.30)	(0.29)
Coupon Rate (fixed, percent)		-0.89***		-3.08***
		(0.18)		(0.79)
Coupon Rate (float, dummy)		-8.69***		-32.23***
		(1.95)		(11.80)
Amortization Profile (payment before maturity, dummy)		2.92*		-14.07
		(1.70)		(8.97)
Duration of Restructuring (years)		17.55**		
		(8.44)		
External Debt Restructuring (dummy)		-2.43		
		(5.47)		
Contant	52.00***	41.02***	78.77***	104.06***
	(0.55)	(9.94)	(2.64)	(6.82)
Episode Fixed Effects	Yes	Yes	Yes	Yes
Number of Countries	15	15	12	12
Number of Restructurings	28	28	16	16
Observations	449	447	111	111
Adjusted R-Squared	0.035	0.262	0.112	0.247

The table reports results of fixed effects OLS regressions. The dependent variable is SZ recovery rate (columns 1 and 2) or Exchange recovery rate (columns 3 and 4) (in %). The main explanatory variable is maturity of instrument at the time of exchange. The control variables are instrument-specific controls and restructuring-specific controls. Significance levels denoted by *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. All regressions include restructuring-specific fixed effects. Robust standard errors in parentheses.

3.2 Bond Prices

Stylized fact 2: The difference in prices between short-term and longer-term bonds decreases in the run-up to the debt exchange.

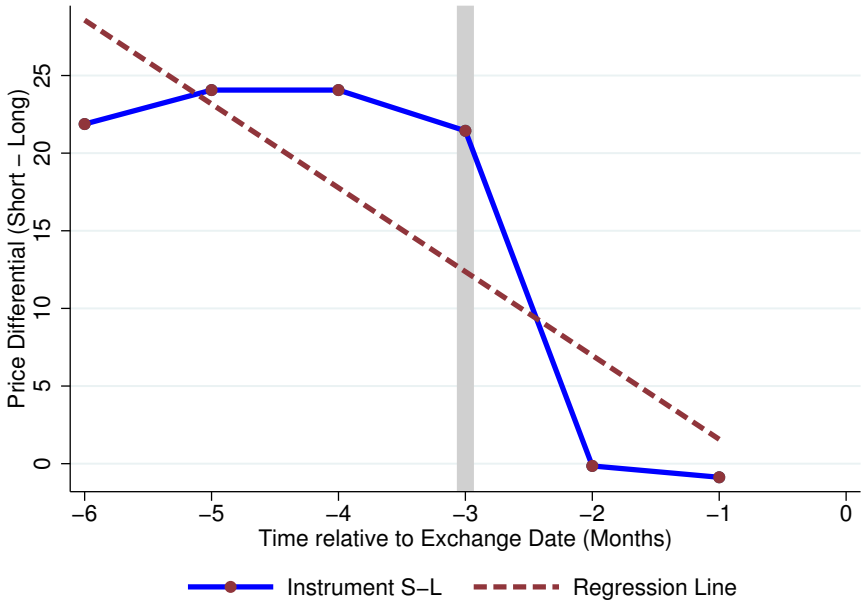
Figure 3 reports bond price differentials—prices of short-term bonds minus prices of longer-term bonds. We present two episodes: (i) Ukraine 2000 external, and (ii) Greece 2011–2012 external. Figure A6 in Appendix D reports two additional episodes: Dominican Republic 2004–2005 external (bonds) and Argentina 2001–2005 external. In all figures, grey bars indicate the date of the announcement of restructuring. Both figures show that prices of short-term and long-term bonds converge in the run-up to the exchange. The observed price dynamics are not specific either to the type of debt (domestic or external) or the restructuring strategy (strictly preemptive, weakly preemptive or post-default).[9,10]

[9] We exclude prices of bonds that are no longer traded on the market.

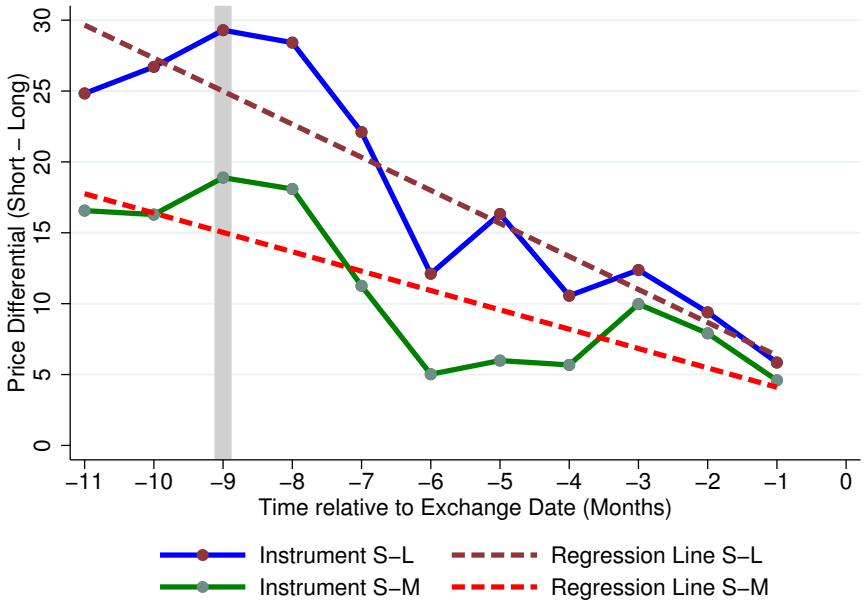
[10] See Guscina et al. (2017) for two episodes of the flattening of the term structure prior to a distress event.

Figure 3: Bond Price Differentials

(i) Ukraine 2000 External



(ii) Greece 2011–2012 External



Note: Grey bars indicate the date of the announcement of restructuring.

Table 3 provides econometric support for this stylized fact. It reports results of panel regressions of the bond price differential on time relative to exchange date and maturity differential with additional instrument- and episode-specific controls:

$$\Delta Price_{j,t}^{S/L} = c + \beta_1 Time_{j,t} + \beta_2 \Delta Maturity_j^{S/L} + \beta_3 y_j + \epsilon_{j,t} \tag{4}$$

where $\Delta Price_{j,t}^{S/L} = Price_{j,t}^{Short} - Price_{j,t}^{Long}$ and $\Delta Maturity_j^{S/L} = Maturity_j^{Short} - Maturity_j^{Long}$ denote bond price differential and maturity differential between short-term bonds and long-term bonds at time t in debt restructuring j respectively. $Time_{j,t}$ denotes the time span until the exchange date (in months) at time t in debt restructuring j. y_j denotes the same vector of episode-specific controls as in Section 3.1.

The main result in Table 3 is that the price of short-term debt approaches the price of longer-term debt from above during the run-up to the exchange. This effect is significant at the 1-percent level, and quantitatively large: over 6 months the price difference diminishes by 9 percentage points (relative to face value of 100). The second result concerns the size of the bond price differential conditional on time: it increases with the maturity differential. For instance, the price of a 1-year bond exceeds the price of a 10-year bond by 4 percentage points. Once we decompose episode-specific fixed effects, we find that restructuring strategy (strictly preemptive vs post-default) and type of debt (external vs domestic) enter as a significant explanatory variable. Moreover, Argentina's 2001–2005 external debt restructuring appears to constitute an outlier in that ceteris paribus, the bond price differential was larger.

The speed of price convergence is approximately the same in the subsample of restructuring episodes where the exchange involves a single menu of new instruments vs different menus, see the results as reported in Table A3 in Appendix D.

Table 3: Panel Regression Results

	Bond Price Differential	Bond Price Differential, with FE	Bond Price Differential, with controls
	(1)	(2)	(3)
	coef/se	coef/se	coef/se
Time Span until the Exchange Date (months)	-1.52***	-1.52***	-1.52***
	(0.08)	(0.08)	(0.08)
Maturity Differential (S - L, year)	-0.49***	-0.49***	-0.49***
	(0.05)	(0.05)	(0.05)
Duration of Restructuring (years)			-10.39
			(7.92)
Argentina 2001–2005 Restructuring (dummy)			43.67**
			(22.12)
Strictly Preemptive Restructuring (dummy)			17.89**
			(7.22)
Weakly Preemptive Restructuring (dummy)			9.79
			(11.71)
External Debt Restructuring (dummy)			16.56**
			(7.07)
Contant	2.88	2.21***	-16.39*
	(3.08)	(0.60)	(8.66)
Episode Fixed Effects	No	Yes	No
Number of Restructurings	12	12	12
Number of Pairs of Instruments	66	66	66
Observations	605	605	605
Adjusted R-Squared	0.224	0.429	0.429

Columns 1 and 3 report random effects OLS regression results and column 2 reports fixed effects OLS regression result. The dependent variable is bond price differential (price of short-term bond − price of long-term bond). The main explanatory variables are the time span until the exchange date (in months) and maturity differential (maturity of short-term bond − maturity of long-term bond) at the time of exchange. The control variables are episode-specific controls. Significance levels denoted by *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Robust standard errors in parentheses.

4 The Model

To identify potential drivers of the observed positive correlation between bond maturity and recovery rate, we rely on a standard asset pricing model adapted to the sovereign debt context. We establish two results. First, prior to the run-up to a debt restructuring prices of short-term bonds exceed prices of longer-term bonds. When a restructuring episode ends with all outstanding instruments being exchanged against the same new instrument (or basket of new instruments), as is typically the case, then this first result implies the main stylized fact that we have documented in Section 3: A higher haircut on short-term bonds. The first result also explains the correlation between haircut and maturity if the haircut is measured following Sturzenegger and Zettelmeyer's approach. Second, in the run-up to a debt restructuring prices of short- and longer-term bonds converge because default risk in the short-term rises. This implies that the extent to which haircuts vary by maturity weakens as the restructuring event approaches. Intuitively, successive bad news in the run-up to the restructuring compress price differentials across maturities and as a consequence, the ex-post haircut differences diminish as well.

4.1 Setup

To derive these results, consider a discrete time environment. In period t, a bond that matures in period m either pays the predefined contractual payment, or the issuer announces a restructuring.[11] In the former case, the bond pays the contractual coupon, c^m, if $t < m$; or the coupon plus principal (unity) if $t = m$. In the latter case, the bond pays a recovery value, ϕ_t^m. The probability of no restructuring between periods t and s, $s > t$, given that no such restructuring occurred until t, equals $\pi_{t,s}$.

The timing of events in each period t is as follows: At the beginning of the period, investors learn whether a restructuring "shock" has occurred, that is whether debt will be restructured and thus pay off ϕ_t^m. They may also receive new information about future restructuring risk that may lead them to update their beliefs about $\pi_{t,s}$. Conditional on the (new) information, the bond pays the contractual payment or the recovery value, and in the former case the ex-coupon price materializes.

To keep the notation simple, we assume that investors are risk neutral and the risk-free gross interest rate is time invariant and equal to β^{-1}. Standard asset pricing then implies that the ex-coupon price equals the expected discounted price and coupon payment in the subsequent period conditional on no restructuring, plus the expected discounted recovery value if a restructuring shock hits in the subsequent period. Formally, letting p_t^m denote the ex-coupon price at date t of a bond maturing at date m,

$$p_t^m = \beta\{\pi_{t,t+1}(c^m + \mathbb{E}_t[p_{t+1}^m|N_{t+1}]) + (1 - \pi_{t,t+1})\,\mathbb{E}_t[\phi_{t+1}^m|A_{t+1}]\} \tag{5}$$

[11] For reasons outside the model (e.g., concerns about financial stability) Treasury bills and trade finances often are treated favorably in debt exchanges. In the empirical analysis, we exclude these instruments.

where N_{t+1} and A_{t+1} denote "no announcement of restructuring in period $t+1$ (and earlier)" and "announcement of restructuring in period $t+1$," respectively. In the last period before maturity, $t = m-1$, the equation holds with $p^m_{t+1} = 1$.

4.2 First Result

Let $\Delta^{s,l}_t \equiv p^s_t - p^l_t$ denote the price difference between a short-term bond with maturity date s and a longer-term bond with maturity date l.[12] With identical coupons and recovery values across bonds, $c^s = c^l$ and $\phi^s_i = \phi^l_i$ for all $i \leq s$, we have

$$\Delta^{s,l}_t = \beta^{s-t}\pi_{t,s}\,\mathbb{E}_t[\Delta^{s,l}_s|N_s] \tag{6}$$

That is, if a positive price difference is expected to prevail at date s—which holds true whenever the longer-term bond is exposed to sufficiently high default risk after period s (or when it has a sufficiently low coupon)—then the current price difference is positive as well. This positive relationship between the current price difference and the expected future price difference only collapses when the coupon or the recovery values of the longer-term bond are sufficiently large relative to those of the short-term bond.

Absent major such differences between the coupons or the recovery values, long-term restructuring risk (and $\mathbb{E}_t[\Delta^{s,l}_s|N_s] > 0$) therefore implies $\Delta^{s,l}_t > 0$: *Short-term bonds are more expensive than longer-term bonds.* This constitutes the first result stated before.

4.3 Second Result

Over time, $\Delta^{s,l}_t$ may change for two fundamentally different reasons. First, because time passes by. And second, because news arrives, in particular about future restructuring risk. With identical coupons and recovery values across bonds, the price difference evolves according to

$$\Delta^{s,l}_{t+1} - \Delta^{s,l}_t = \beta^{s-t-1}\pi_{t+1,s|t+1}\,\mathbb{E}_{t+1}[\Delta^{s,l}_s|N_s] - \beta^{s-t}\pi_{t,s|t}\,\mathbb{E}_t[\Delta^{s,l}_s|N_s] \tag{7}$$

where the subscript of the probability of no restructuring now also indicates the information set.[13]

Absent new information, the change in the price difference reduces to

$$\Delta^{s,l}_{t+1} - \Delta^{s,l}_t = \beta^{s-t-1}\pi_{t+1,s|t+1}\,\mathbb{E}_t[\Delta^{s,l}_s|N_s](1-\beta\pi_{t,t+1|t}) \tag{8}$$

[12] This price difference is given by

$$\begin{aligned}\Delta^{s,l}_t &\equiv \beta\{\pi_{t,t+1}(c^s - c^l + \mathbb{E}_t[\Delta^{s,l}_{t+1}|N_{t+1}]) + (1-\pi_{t,t+1})\,\mathbb{E}_t[\phi^s_{t+1} - \phi^l_{t+1}|A_{t+1}]\}\\ &= (c^s - c^l)\sum_{i=t+1}^{s}\beta^{i-t}\pi_{t,i} + \sum_{i=t+1}^{s}\beta^{i-t}\pi_{t,i-1}(1-\pi_{i-1,i})\,\mathbb{E}_t[\phi^s_i - \phi^l_i|A_i] + \beta^{s-t}\pi_{t,s}\,\mathbb{E}_t[\Delta^{s,l}_s|N_s]\end{aligned}$$

[13] That is, $\pi_{t+1,s|t}$ denotes the probability of no restructuring between periods $t+1$ and s conditional on information at time t.

which is positive under the assumptions leading to the first result discussed above. That is, *absent new information prices further diverge.*[14]

If new information implies sufficiently increased default risk in the short term, however, then $\Delta^{s,l}$ may fall over time. Intuitively, an increase of short-term default risk proportionally lowers the prices of short- and longer-term debt. In absolute terms, the *price of short-term debt thus falls more strongly.*[15] This constitutes the second result. It reflects the fact that lower expected payoffs support lower contemporaneous prices. Note that the result continues to hold if new information about recovery values arrives as long as the recovery values are identical for bonds with different maturities.

4.4 Interpretation

By selecting episodes leading to restructuring we implicitly select episodes characterized by rising short-term default risk. Ceteris paribus, the model predicts the price difference $\Delta^{s,l}$ to be positive during such episodes but decreasing over time. The positive price difference explains the empirical finding of higher haircuts on shorter-term debt. The decreasing price difference explains the empirical finding that the relation between maturity and haircut is weaker when the restructuring is closer.

5 Default Probability Term Structure

We have argued in Section 4 that basic theory explains the two stylized facts—higher haircuts on short-term bonds and smaller haircut differences with shorter window length—if short-term default risk rises in the run-up to a restructuring. We now investigate empirically the relation between bond prices and short-term default risk. In line with theory, we document an additional stylized fact: in the run-up to an exchange the term structure of default risk becomes elevated at the short end.

Stylized fact 3: The term structure of default probability becomes elevated at the short end in the run-up to the exchange.

Using the method introduced in Ranciere (2002), we estimate the term structure of default risk. In our estimation exercise, the main inputs are (a) the Credit Default Swap (hereafter CDS) spread curve, and (b) the US Treasury bill yield curve proxying for the risk-free interest curve. Following Ranciere (2002), we assume (i) no arbitrage, (ii) risk-neutrality, and (iii) the exchange involving a single menu of new instruments for all restructured instruments.

Under these assumptions, we use annual yield curve data and linear interpolations to construct estimates of the CDS spread curve for horizons between 1 and 120 months, at the monthly frequency. Based on this, we use the no arbitrage condition to derive the forward default spread

[14] Deterministic price divergence may also result due to different coupons or expected recovery values ($c^s < c^l$ and $\phi^s < \phi^l$).

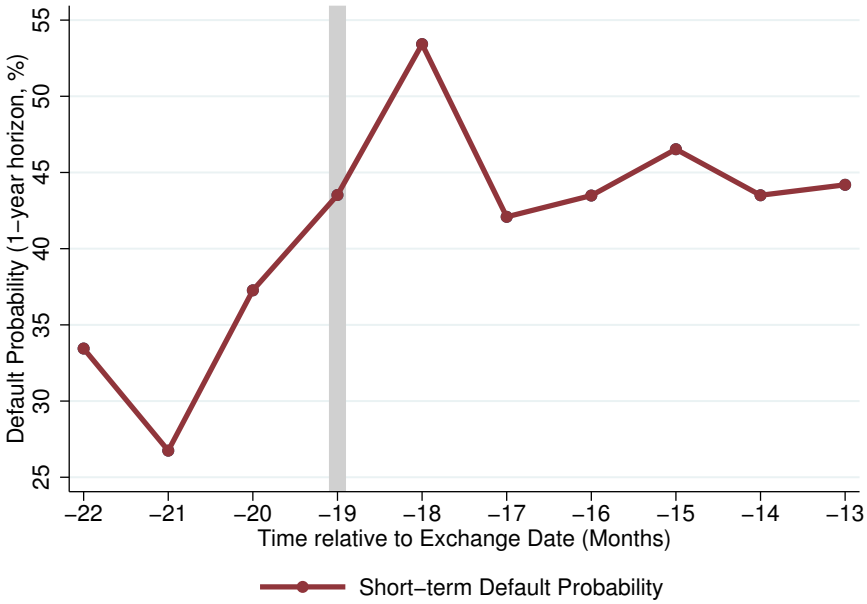
[15] The price difference might also fall if the expected future price difference is updated downward.

curve. This curve measures the implied one-month ahead conditional default risk. Finally, we use the risk-neutrality assumption and the fact that the exchange involves a single menu of new instruments for all restructured instruments to derive the implied one-month ahead conditional default probability. Appendix E discusses the details.

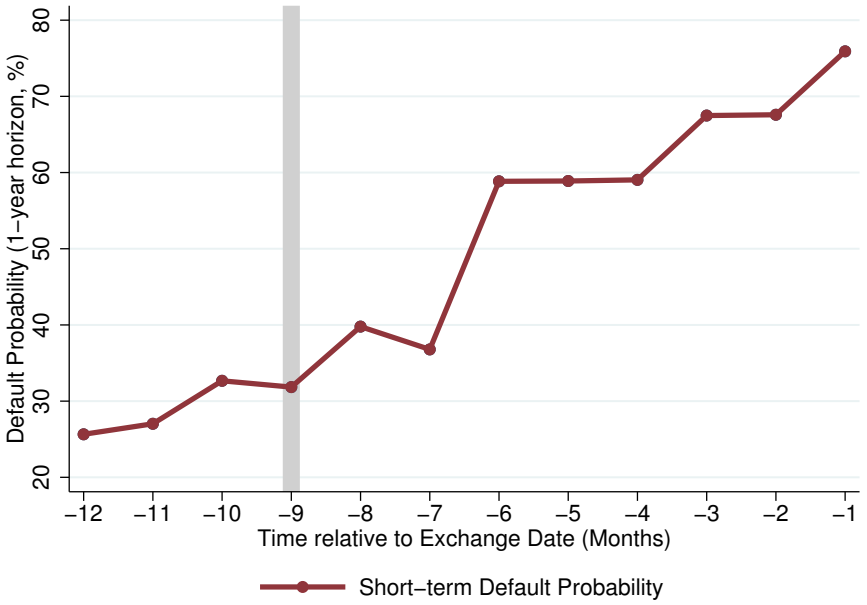
Figure 4 reports our conditional default probability estimates at the 1-year horizon. We present two cases: (i) Ecuador 1999–2000 external and (ii) Greece 2011–2012 external. Figure A8 in Appendix F reports the case of Argentina 2001–2005 external. In all figures, grey bars indicate the date of the restructuring announcement. Both figures show that the conditional short-term default probability (1 year) increases in the run-up to the exchange. This observed rise is not specific to restructuring strategy (strictly preemptive, weakly preemptive or post-default) and we also observe it at the 2-year and 3-year horizon, as reported in Table A4 in Appendix F.

Figure 4: Default Probability Term Structure

(i) Ecuador 1999–2000 External



(ii) Greece 2011–2012 External



Note: Grey bars indicate the date of the announcement of restructuring.

Table 4: Panel Regression Results

	Default Probability 1–year Horizon(%)	Default Probability 1–Year Horizon, with FE	Default Probability 1–Year Horizon, with controls
	(1)	(2)	(3)
	coef/se	coef/se	coef/se
Time Span until the Exchange Date (months)	1.09***	2,70***	2.70***
	(0.26)	(0.37)	(0.37)
Duration of Restructuring (years)	-3.11**		-15.68***
	(1.43)		(2.66)
Strictly Preemptive Restructuring (dummy)			-37.98***
			(7.37)
Contant	61.20***	71.92***	115.36***
	(4.61)	(4.06)	(11.06)
Episode Fixed Effects	No	Yes	No
Number of Restructurings	3	3	3
Number of Time Periods (months)	12	12	12
Observations	36	36	36
Adjusted R-Squared	0.373	0.628	0.628

Columns 1 and 3 report random effects OLS regression results and column 2 reports a fixed effects OLS regression result. The dependent variable is the conditional default probability at the 1-year horizon (%). The main explanatory variable is the time span until the exchange date (in months). The control variables are episode-specific controls. Significance levels denoted by *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Robust standard errors in parentheses.

Table 4 provides econometric support for the stylized fact. It presents results of panel regressions of the conditional default probability at the 1-year horizon on the time span until the exchange date with additional episode-specific controls:

$$Default_{j,t} = c + \beta_1 Time_{j,t} + \beta_2 y_j + \epsilon_{j,t} \tag{9}$$

where $Default_{j,t}$ denotes the conditional default probability over the 1-year horizon at time t in debt restructuring j. $Time_{j,t}$ denotes the time span until the exchange at time t in debt restructuring j. y_j denotes the same vector of episode-specific controls as in Section 3.1.

The main result in Table 4 is that the conditional default probability increases in the run-up to the exchange. This effect is significant at the 1-percent level, and quantitatively large: over 6 months the default risk increases by 7–16 percentage points. Once we decompose episode-specific fixed effects, we find that restructuring strategy (strictly preemptive vs post-default) enters as a significant explanatory variable. Duration of restructuring turns out to significantly reduce the conditional default probability possibly because default probability gradually increases for restructurings with longer duration.

6 Conclusion

This paper studies creditor losses (haircuts) across sovereign debt instruments during debt restructurings. We show that, contrary to conventional wisdom, creditors are not treated equally in sovereign debt exchanges. We first construct a comprehensive dataset on instrument-specific haircuts in recent sovereign debt restructuring episodes affecting private creditors. We demonstrate that haircuts on shorter-term debt tend to be larger than those on longer-term debt, and propose a simple asset pricing model that can theoretically explain this fact. We also show that the data supports the causal link among short-term default risk, bond prices and maturity-specific haircuts suggested by the model.

Our findings highlight a regularity that has not received much attention so far and has potential implications for important dimensions of debt policy including, pari passu and its limitations; the negotiation process during debt restructuring episodes; the value at risk of debt securities of different maturities (and thus, of financial institutions that hold these instruments); or the optimal debt management more broadly. However, further research is required before drawing normative conclusions or policy prescriptions, as the findings in this paper do not allow to conclude that specific bond maturities are better or worse than others for issuers or for investors.

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Appendix A Dataset: Selected Recent Restructurings

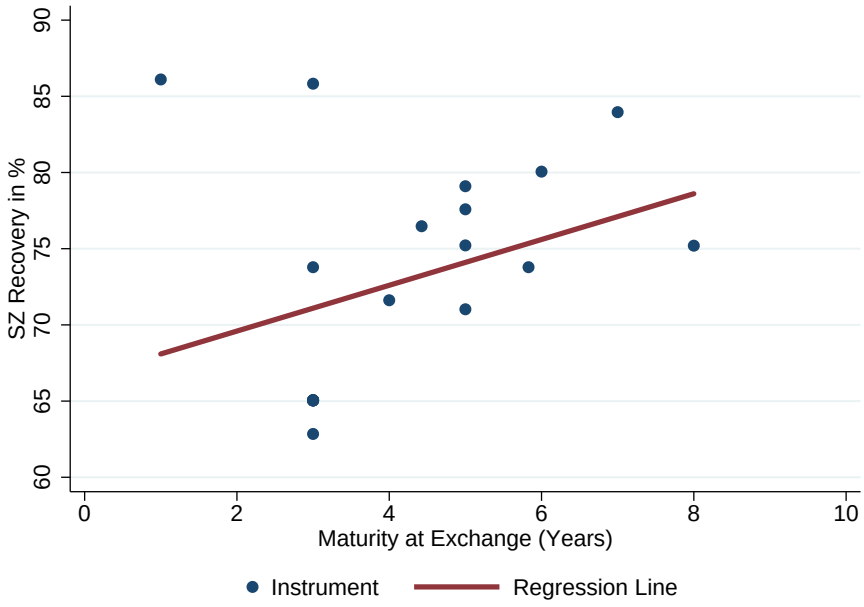
Table A1: Selected Recent Sovereign Debt Restructurings (1998–2015)

Episode	**Announcement of Default** (Start)	**Exchange** (End)	**Restructuring Strategy**	**SZ Haircut Differential** (S- and L-term, percent)	**Maturity of Short- and Long-term Debt** (years)
External					
Pakistan	Jan-99	Dec-99	Strictly Preemptive	25.7	0 and 2.5
Russia (PRINs & IANs)	Nov-98	Aug-00	Post-Default	19.2	-0.1 and 10.9
Ecuador	Jan-99	Aug-00	Post-Default	28.2	1.7 and 24.5
Uruguay	Mar-03	May-03	Strictly Preemptive	8.4	0.5 and 24.2
Argentina (Global Exchange)	Nov-01	Jun-05	Post-Default	13.4	0.1 and 22.6
Grenada	Oct-04	Nov-05	Weakly Preemptive	21.4	0 and 13.0
Belize	Aug-06	Feb-07	Weakly Preemptive	22.4	0 and 8.0
Greece	Jul-11	Mar-12	Strictly Preempitve	45.4	0 and 28.5
Saint Kitts and Nevis	Jun-11	Apr-12	Weakly Preemptive	10.3	0 and 8
Domestic					
Russia (GKOs resident)	Aug-98	May-99	Post-Default	2.4	-0.3 and 0.3
Argentina (Phrase I)	Oct-01	Nov-01	Strictly Preemptive	7.8	0.5 and 29.6
Uruguay	Mar-03	May-03	Strictly Preemptive	10.4	0.0 and 8.8
Jamaica	Feb-13	Mar-13	Strictly Preemptive	0.9	0 and 33.2
Cyprus	May-13	Jul-13	Strictly Preemptive	14.7	1 and 2.7

Appendix B Haircuts/Recovery Rates

Figure A1: SZ Recovery and Maturity

(i) Belize 2006–2007 External



(ii) Saint Kitts and Nevis 2011–2012 Domestic

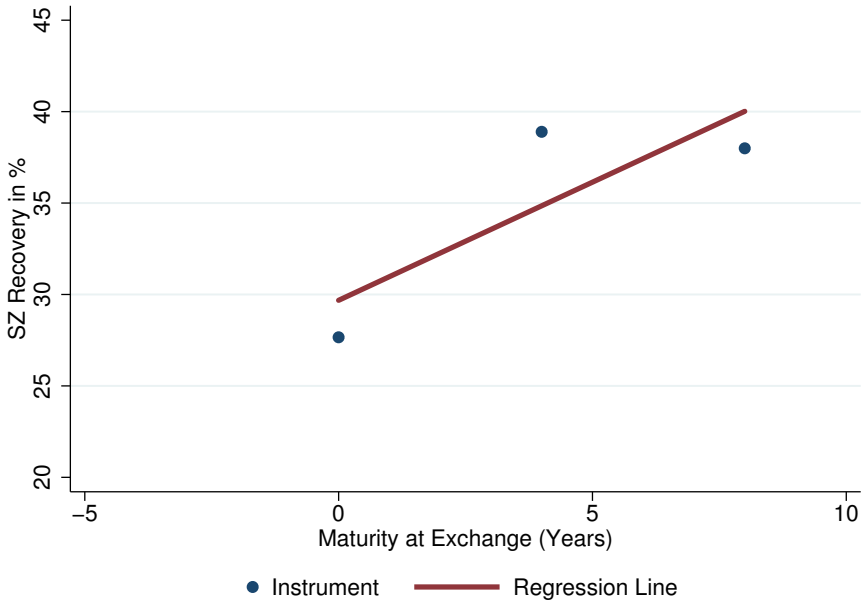
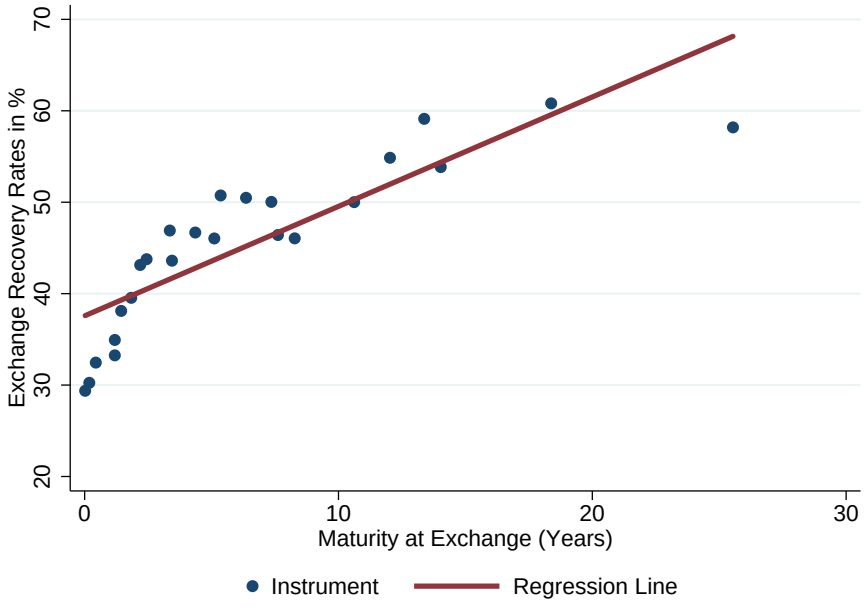


Figure A2: Exchange Recovery and Maturity

(i) Greece 2011–2012 External



(ii) Urugauy 2003 Domestic

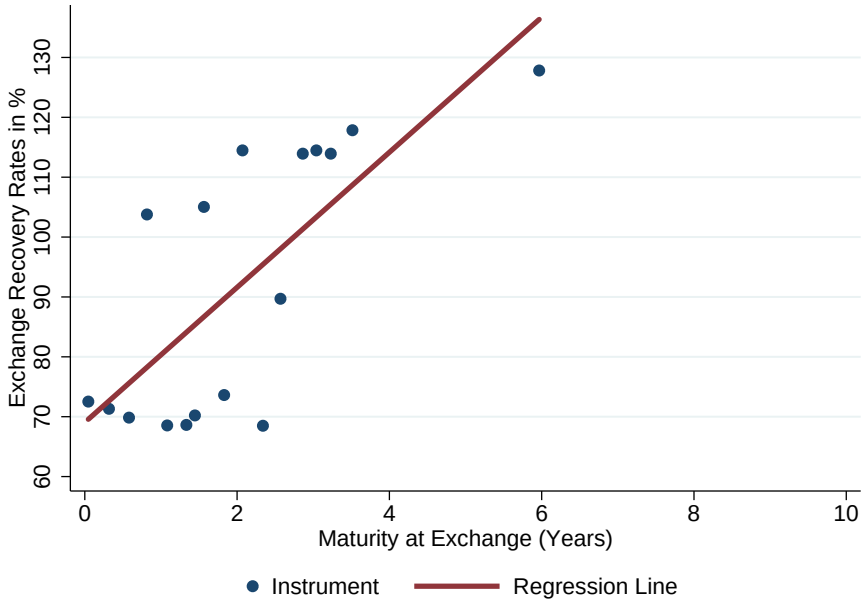
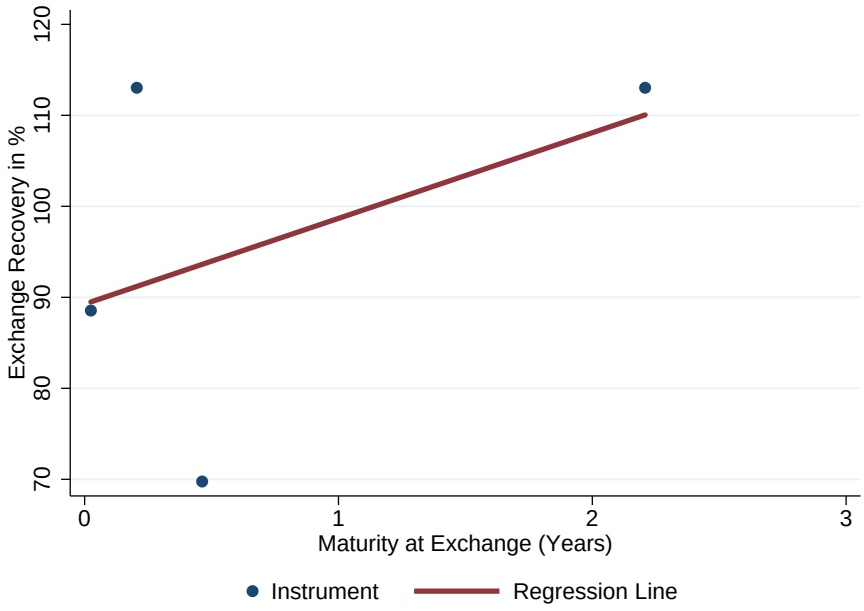
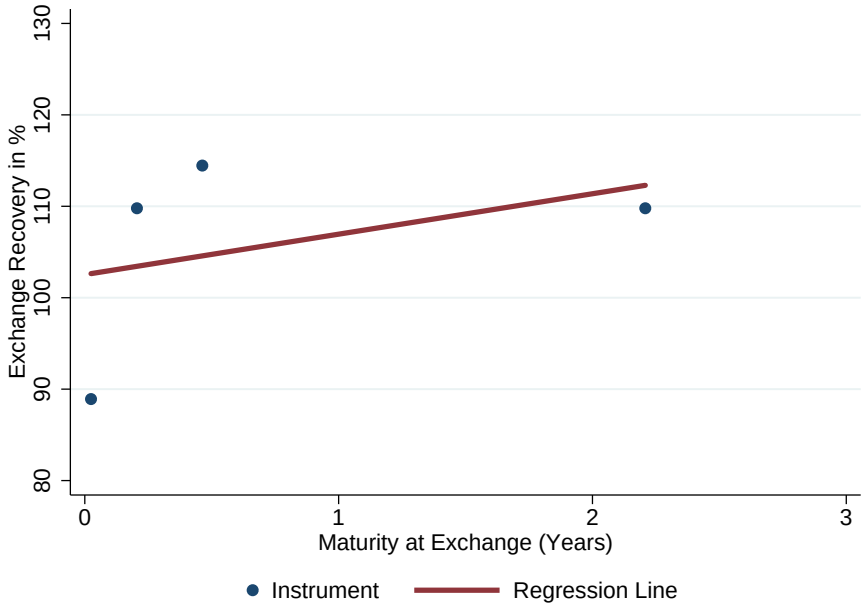


Figure A3: Exchange Recovery and Maturity at Different Points of Time

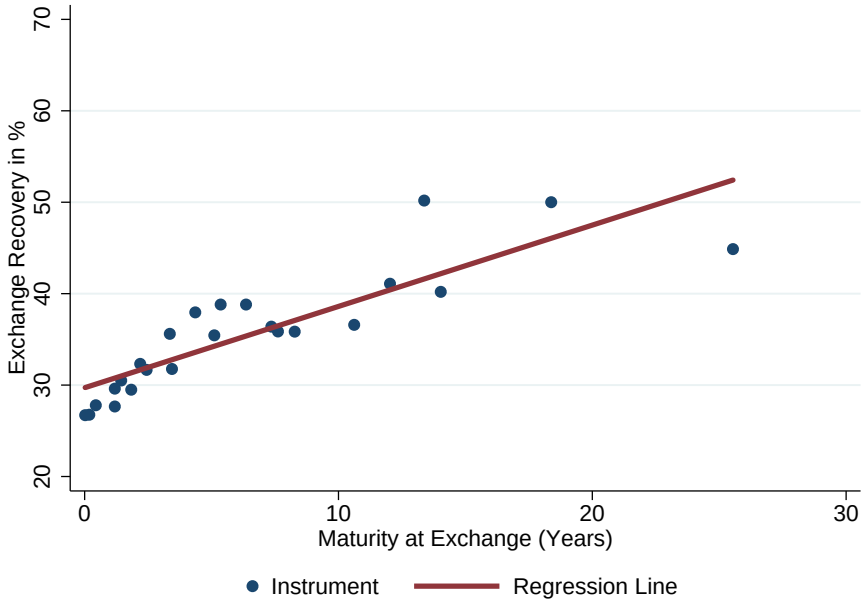
(i) Pakistan 1999 External — 6 months before the announcement



(ii) Pakistan 1999 External — 6 months after the announcement



(iii) Greece 2011–2012 External — 6 months before the announcement



(iv) Greece 2011–2012 External — 6 months after the announcement

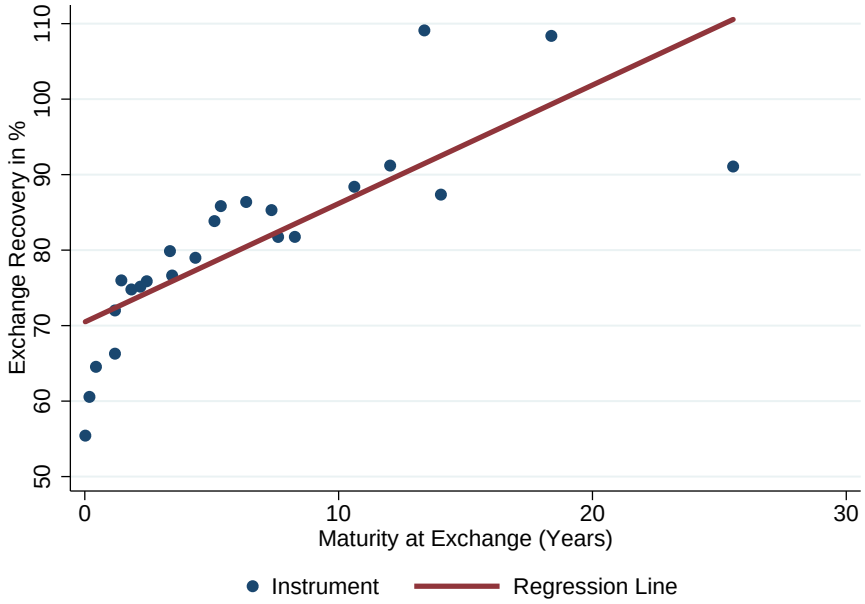


Table A2: Cross-sectional Regression Results

	Exchange Recovery Rate (6 months before announcement)	Exchange Recovery Rate, with controls	**Exchange Recovery Rate (6 months after (announcement)**	Exchange Recovery Rate, with controls
	(1)	(2)	(3)	(4)
	coef/se	coef/se	coef/se	coef/se
Maturity of Instrument (years)	0.54**	0.54**	5.74***	5.79***
	(0.24)	(0.24)	(1.74)	(1.81)
Coupon Rate (fixed, percent)		-0.76		0.86
		(0.64)		(4.95)
Coupon Rate (float, dummy)		-11.31		4.53
		(9.18)		(74.22)
Amortization Profile (payment before maturity, dummy)		-13.61*		33.03
		(7.12)		(56.42)
Duration of Restructuring (years)				
External Debt Restructuring (dummy)				
Contant	68.70***	76.68***	80.04***	71.25
	(1.89)	(5.64)	(15.31)	(42.88)
Episode Fixed Effects	Yes	Yes	Yes	Yes
Number of Countries	12	12	12	12
Number of Restructurings	16	16	16	16
Observations	104	104	111	111
Adjusted R-Squared	0.054	0.118	0.104	0.108

The table reports results of fixed effects OLS regressions. The dependent variable is Exchange recovery rate (in %). The main explantory variable is maturity of instrument at the time of exchange. Columns 1 and 2 report results for Exchange recovery rate 6 months before the announcement and columns 3 and 4 report results for Exchange recovery rate 6 months after the announcement. The control variables are instrument-specific controls and restructuring-specific controls. Significance levels denoted by *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. All regressions include restructuring-specific fixed effects. Robust standard errors in parentheses.

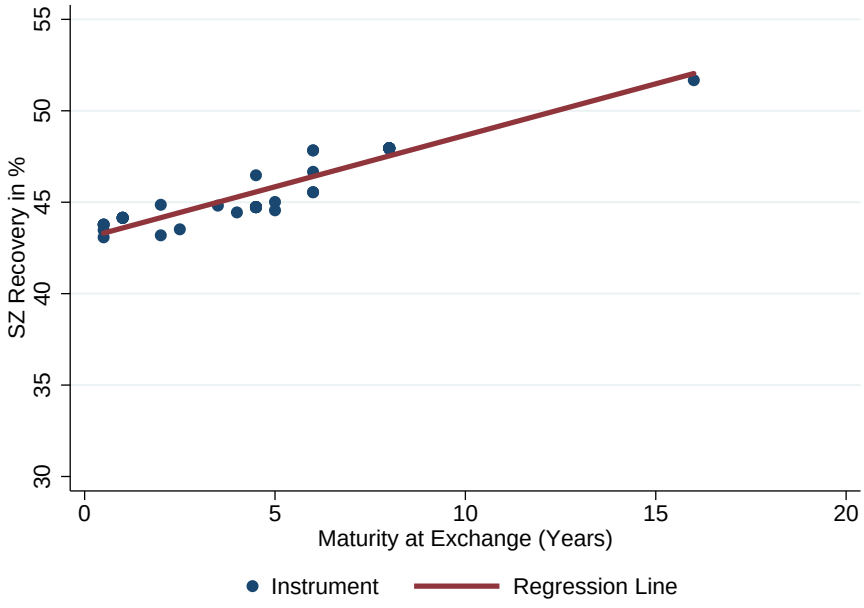
Appendix C SZ Haircuts Robustness Check

One advantage of the SZ haircut measure is that it is relatively robust to the posited exit yield of new instruments because both the cash flows of old and new instruments are discounted at the same yield. However, there still remains the question when the exit yield should be measured. To address this concern, we conduct a robustness check and recompute haircuts for higher (exit yield +1 percent) and lower (exit yield -1 percent) yields.

Figure A4 and A5 report results for Dominica 2003–2004 domestic and Belize 2006–2007 external, respectively. In both cases, SZ recovery rates on short-term debt are lower than those on long-term debt whichever yield is used.

Figure A4: SZ Recovery Rate

(i) Dominica 2003–2004 Domestic, exit yield +1 percent



(ii) Dominica 2003–2004 Domestic, exit yield -1 percent

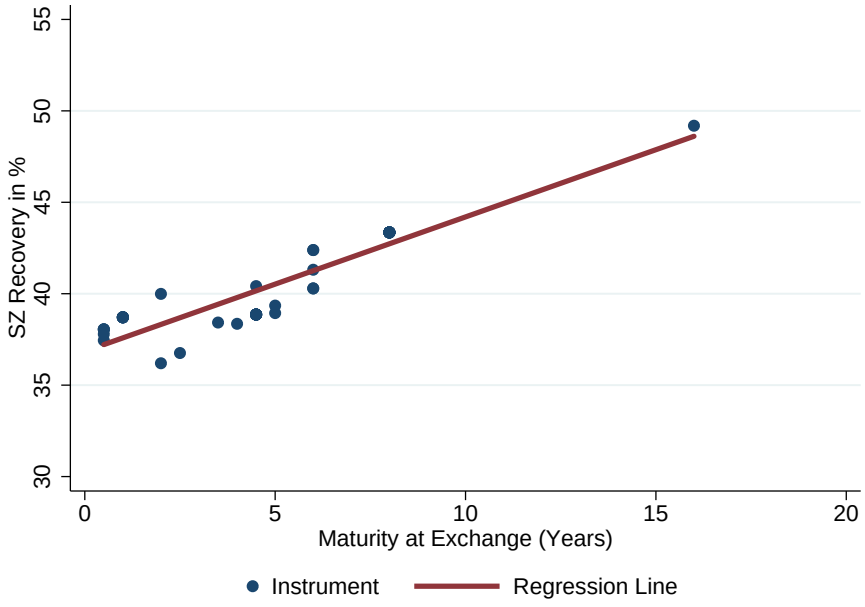
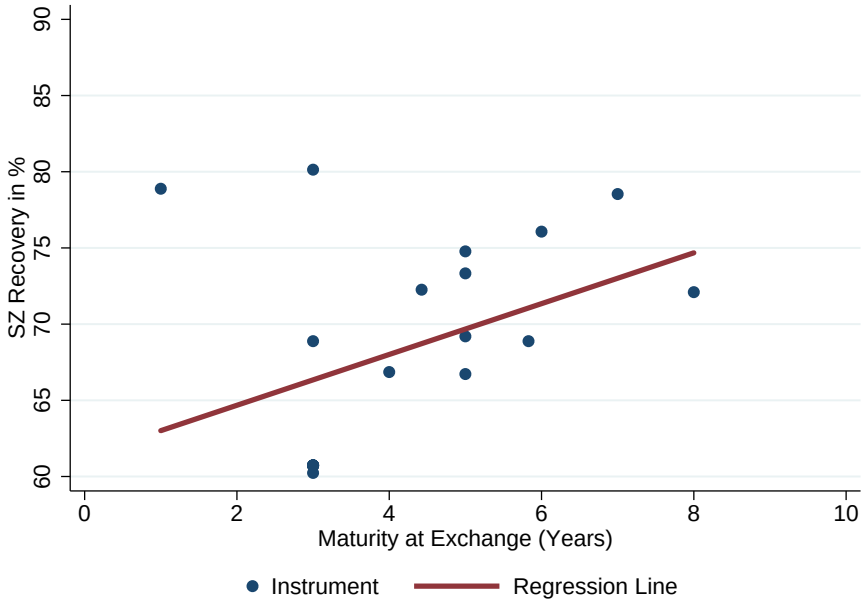
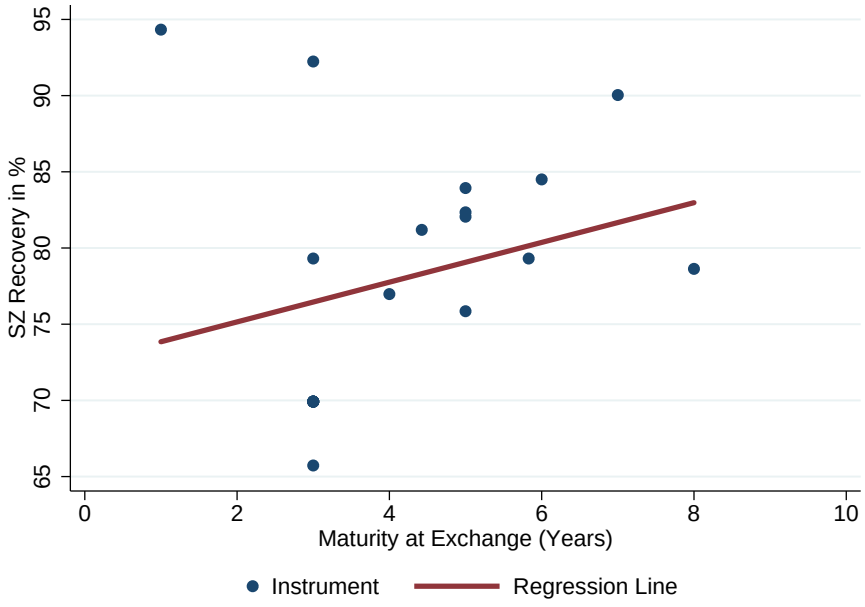


Figure A5: SZ Recovery Rate

(i) Belize 2006–2007 External, exit yield +1 percent



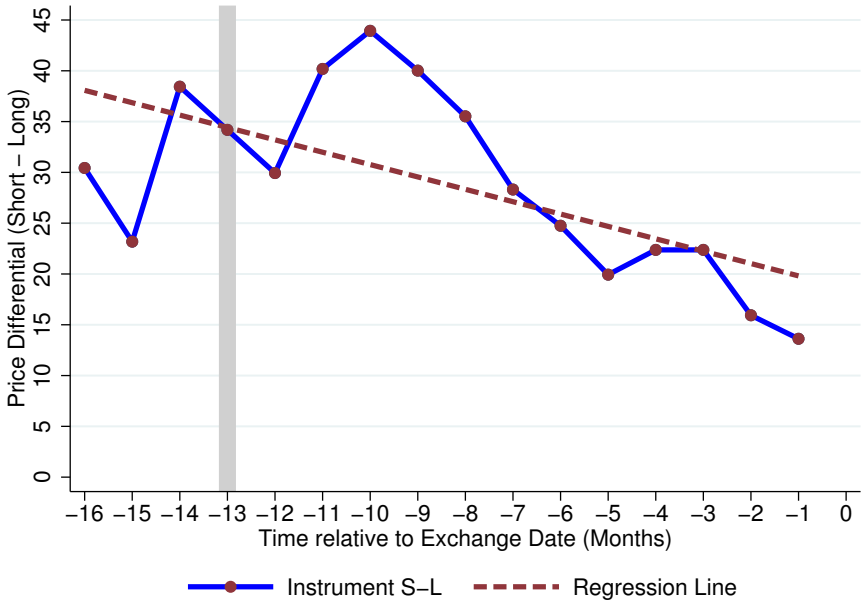
(ii) Belize 2006–2007 External, exit yield -1 percent



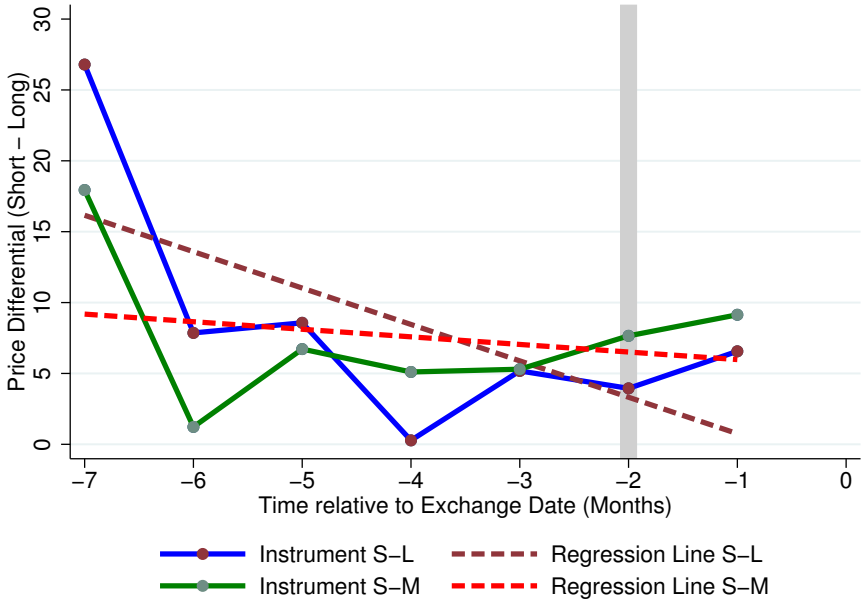
Appendix D Bond Prices

Figure A6: Bond Price Differentials

(i) Dominican Republic 2004–2005 External (bonds)



(ii) Argentina 2001–2005 External



Note: Grey bars indicate the date of the announcement of restructuring.

Table A3: Panel Regression Results

	Bond Price Differential, Symmetric	Bond Price Differential, Symmetric, with controls	**Bond Price Differential, Non-Symmetric**	Bond Price Differential, Non-Symmetric, with controls
	(1)	(2)	(3)	(4)
	coef/se	coef/se	coef/se	coef/se
Time Span until Exchange Date (months)	-1.42***	-1.43***	-1.79***	-1.81***
	(0.09)	(0.10)	(0.15)	(0.15)
Maturity Differential (S - L, year)	-0.51***	-0.40***	0.34	0.81***
	(0.05)	(0.05)	(0.24)	(0.22)
Duration of Restructuring (years)		-37.95***		-29.37***
		(2.41)		(4.25)
Argentina 2001–2005 restructuring (dummy)				
Strictly Preemptive Restructuring (dummy)		-114.63***		19.75***
		(7.23)		(2.02)
Weakly Preemptive Restructuring (dummy)				-15.74***
				(5.98)
External Debt Restructuring (dummy)		17.43***		48.69***
		(1.56)		(5.11)
Contant	7.81**	127.09***	1.05	-18.02***
	(3.70)	(7.72)	(5.46)	(2.74)
Episode Fixed Effects	No	No	No	No
Number of Restructuirngs	6	6	6	6
Number of Pairs of Instruments	51	51	15	15
Observations	487	487	118	118
Adjusted R-Squared	0.190	0.403	0.289	0.787

The table reports results of random effects OLS regressions. The dependent variable is bond price differential (price of short-term bond - price of long-term bond). The main explanatory variables are the time span until the exchange date (in months) and maturity differential (maturity of short-term bond - maturity of long-term bond) evaluated at the time of exchange. Columns 1 and 2 report regression results for restructurings with exchanges involving a single menu of new instruments. Columns 3 and 4 report regression results for restructurings with exchanges involving different menus of new instruments. The control variables are episode-specific controls. Significance levels denoted by *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Robust standard errors in parentheses.

Appendix E Estimation of Term Structure of Default Risk

We follow Ranciere (2002).

1st Step: Constructing the monthly yield curve

We interpolate the annual yield curve linearly for both (i) CDS spreads and (ii) risk-free interest rates as follows:

$$CDS_t^{M,j} = CDS_t^{A,i} + \frac{j-i}{12}(CSD_t^{A,i+12} - CDS_t^{A,i})$$
$$RF_t^{M,j} = RF_t^{A,i} + \frac{j-i}{12}(RF_t^{A,i+12} - RF_t^{A,i}) \tag{A1}$$
$$\text{for } i = 0, 12, 23, 36, ..., 108 \text{ and } j = i+1, i+2, ..., i+11$$

The annual yield for 0-year horizon is assumed to be zero; $CDS_t^{A,0} = RF_t^{A,0} = 0$.

where $CDS_t^{M,j}$ and $RF_t^{M,j}$ denote estimated monthly CDS spreads and risk-free interest rates, respectively. $CDS_t^{A,i}$ and $RF_t^{A,j}$ denote observed annual CDS spreads and risk-free interest rates, respectively.

2nd Step: Computing a forward default spread curve

We apply the no arbitrage condition to derive the forward default spread curve (in month 1,2,3,...., 120) as follows:

$$(1 + RF_t^{M,h,j} + CDS_t^{M,h,j}) = (1 + RF_t^{M,j} + CDS_t^{M,j})/(1 + RF_t^{M,h} + CDS_t^{M,h}) \tag{A2}$$
$$\text{for } i = 1, 2, 3, ..., 120 \text{ and } h = j - 1$$

where $CDS_t^{M,h,j}$ and $RF_t^{M,h,j}$ denote estimated forward default spreads and forward risk-free interest rates, respectively. The forward default spread measures the implied one-period ahead conditional default risk.

3rd Step: Computing the term structure of conditional default probabilities

We use risk-neutrality and the fact that the exchange involves a single menu of new instruments for all restructured instruments to derive the implied one-period ahead conditional default risk:

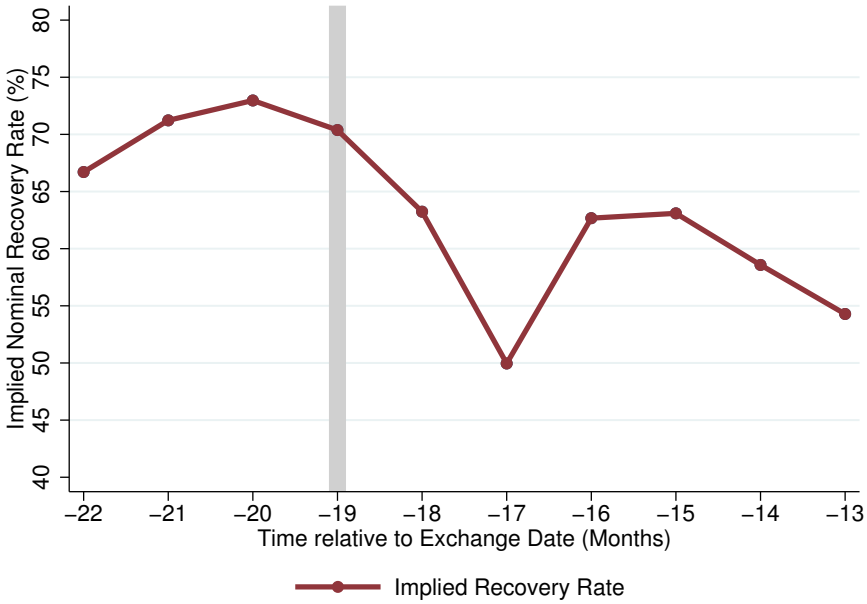
$$(1 + RF_t^{M,h,j}) = (1 - \hat{P}_t^{h,j}) * (1 + RF_t^{M,h,j} + CDS_t^{M,h,j}) + \hat{P}_t^{h,j} * \hat{R}_t$$
$$s.t. \quad \hat{R}_t = argmin \frac{1}{n} \sum_{i=1}^{n} \{p_t^i - \hat{p}_t^i(\{RF_t^{M,k,k+1}\}_{k=h}^{H-1}, \hat{P}_t^{j,h}, \hat{R}_t)\}^2 \tag{A3}$$
$$\text{for } j = 1, 2, 3, ..., 120 \text{ and } h = j - 1$$

where $\hat{P}_t^{h,j}$ and $\hat{R}_t$ are implied one-period ahead conditional default probability and time variant recovery. The implied time-variant recovery is estimated to minimize mean squared deviations of the estimated bond prices $\hat{p}_t^i(\{RF_t^{M,k,k+1}\}_{k=h}^{H-1}, \hat{P}_t^{j,h}, \hat{R}_t)$—based on inputs of streams of risk-free interest rates until maturity, estimated conditional default risk and time-variant recovery—from the observed price p_t^i. This implied time-variant recovery converges to the observed nominal recovery at the time of exchange as reported in Figure A7.[16]

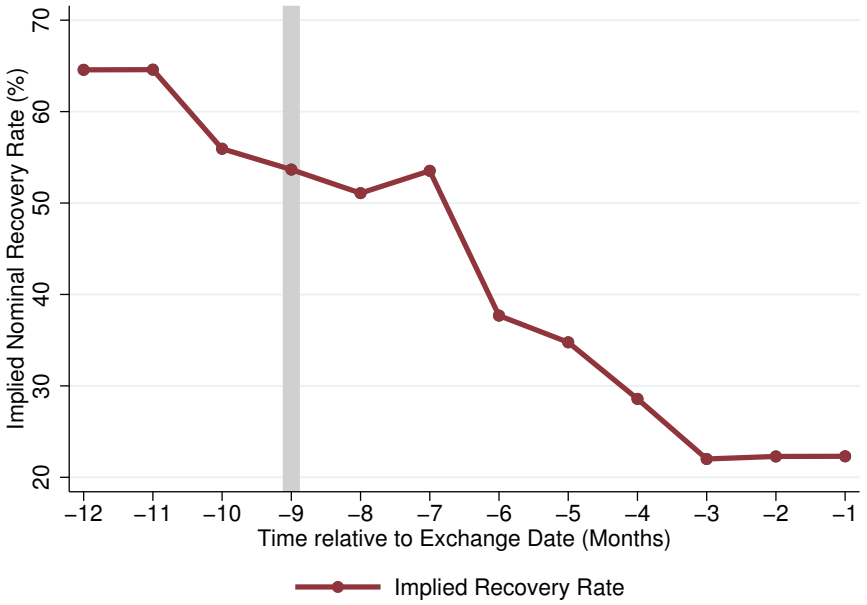
[16] JP Morgan (2000) proposes an alternative approach to compute implied recovery rates.

Figure A7: Implied Nominal Recovery Rate

(i) Ecuador 1999–2000 External



(ii) Greece 2011–2012 External



Note: Grey bars indicate the date of the announcement of restructuring.

Appendix F Default Probability Term Structure

Figure A8: Default Probability Term Structure

(i) Argentina 2001–2005 External

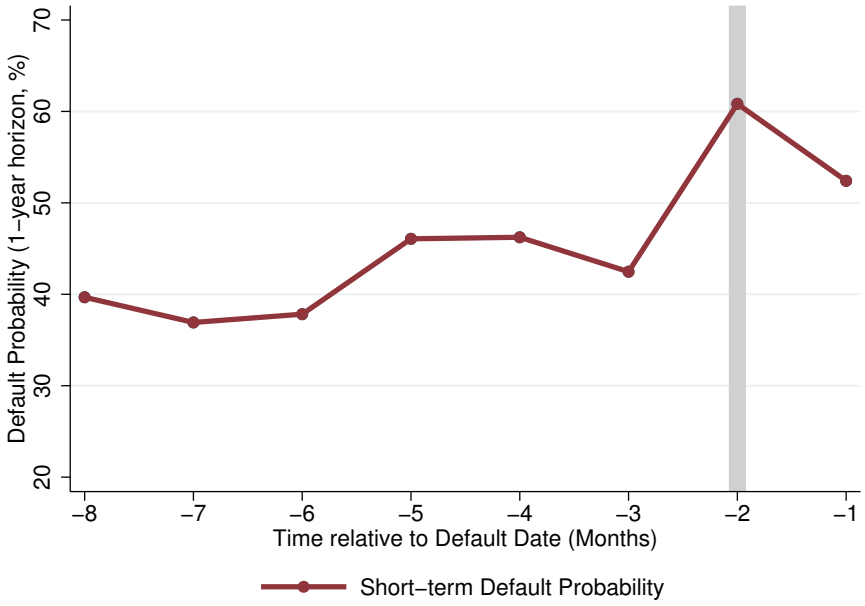


Table A4: Panel Regression Results

	Default Probability 2-Year Horizon (%)	Default Probability 2-Year Horizon, with controls	**Default Probability 3-Year Horizon (%)**	Default Probability 3-Year Horizon, with controls
	(1)	(2)	(3)	(4)
	coef/se	coef/se	coef/se	coef/se
Time Span until the Exchange Date (months)	0.87***	2.16***	0.68***	1.87***
	(0.22)	(0.31)	(0.21)	(0.30)
Duration of Restructuring (years)	-1.86	-12.01***	-1.40	-10.72***
	(1.19)	(2.23)	(1.14)	(2.19)
Strictly Preemptive Restructuring (dummy)		-30.69***		-28.20***
		(6.17)		(6.06)
Contant	52.56***	96.28***	48.28***	88.45***
	(3.83)	(9.27)	(3.65)	(9.09)
Episode Fixed Effects	No	No	No	No
Number of Restructurings	3	3	3	3
Number of Time Periods (months)	12	12	12	12
Observations	36	36	36	36
Adjusted R-Squared	0.336	0.606	0.252	0.543

The table reports results from random effects OLS regressions. The dependent variable is the conditional default probability at the 2-year and 3-year horizon (%). The main explanatory variable is the time span until the exchange date (in months). Columns 1 and 2 report regression results for the default probability at the 2-year horizon and columns 3 and 4 report regression results for the default probability at the 3-year horizon. The control variables are episode-specific controls. Significance levels denoted by *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$. Robust standard errors in parentheses.