Evidence for Dynamic Contracts in Sovereign Bank Lending *

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Abstract

This paper presents direct evidence for self-enforcing dynamic contracts in sovereign bank lending. Unlike the existing empirical literature, its instrumental variables method allows for distinguishing a direct influence of past repayment problems on current spreads (a "punishment" effect in prices) from an indirect effect through higher expected future default probabilities. Such a punishment provides positive surplus to lenders after a default, a feature that characterizes dynamic contracts. Using data on bank loans to developing countries between 1973-1981 and constructing continuous variables for credit history, we find evidence that most of the influence of past repayment problems is through the direct, punishment channel.

Key Words: reputation, dynamic contracts, sovereign bank loan spreads, rational expectations, default risk.

JEL Classification: C73, D86, F34, G12, G14, G15.

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1 Introduction

A central issue in limited contract enforceability is that if the identity of a defaulting party is forgotten, then it can reenter the market with a fresh start. The acts of countries and major banks, however, are remembered, thus they cannot simply walk away after breaking their commitments. This implies that one can view sovereign bank lending as a long-term relationship (a dynamic contract) between borrowers and lenders. In such contracts, parties honor their obligations in order to influence the terms of future interactions. This can mean that a default or any other form of misbehavior leads to an increase in borrowing costs or, as an extreme form, a capital market exclusion. As shown by Kletzer and Wright (2000), Wright (2002), and Kovrijnyikh and Szentes (2007), such punishments are fully compatible with competitive markets, due to the repeated lender interactions that characterize dynamic contracts.

There is indeed evidence that a default leads to an increase in spreads (Ozler, 1993, Eichengreen and Mody, 1999, Reinhart et al, 2003). This literature, however, has been unable to identify the precise mechanism of this effect. In particular, it cannot tell apart a dynamic contract argument from a signaling alternative: a default reveals some adverse information about the expectation of the debtor's future output (for example, the type of the debtor), which hurts its future outcomes (Eaton, 1996, Sandleris, 2008). This mechanism can be labeled as "domestic costs" or "reputational spillovers" in the classification of Panizza et al (2009).

Can we separate the two mechanisms? Our identification strategy is based on the following observation. In the signaling case, reputational spillovers indirectly contribute to a higher spread by increasing the probability of a future default, but lenders still earn zero expected profits. In contrast, in dynamic contracts punishments are incorporated directly into future prices, giving positive surpluses to lenders. The existing empirical literature cannot determine whether a default is followed by an increase in sovereign bank spreads in excess of the increase of future nonregiment risk. We present empirical evidence that a sovereign default is followed by positive lender surpluses, which is consistent with the dynamic contract mechanism.

Such evidence has immediate consequences for understanding sovereign risk, as it points to the presence of dynamic contracts as a repayment mechanism. It is also relevant for the broad context of repeated games, relational contracts and reputations: besides prices and markets, "relationships provide an alternative mechanism that also plays an important role in allocating resources " (Samuelson, 2006, Section 1.3). In both cases, there is very little direct evidence on dynamic incentives themselves - the sovereign risk literature, for example, usually calibrates its models to match various aggregate outcomes, like the cyclicality of sovereign debt flows and interest rate spreads, or the timing and frequency of default.

The purpose of our paper is to contribute to the empirical sovereign risk literature in two major respects. First, and most importantly, our methodology is able to distinguish between any direct effect of a bad repayment history on the spread, and the indirect one going through increased default probability. Second, we deploy different econometric techniques and variables in order to control for country fixed effects, a problem that is important in the framework of this analysis and has not been thoroughly dealt with in most of the previous research.

The data on sovereign bank loan spreads is from the World Bank's publication "Borrowing in International Capital Markets" for the period 1973-1981, on 37 developing countries. This period was the heyday of syndicated bank lending to sovereigns. While most of the variables utilized in the paper are those suggested by the literature, we create continuous measures of past and future default, which are based on arrears data from the World Bank's Global Development Finance. These variables are compatible with country fixed effects.

The main estimation strategy used in the paper is a structural-form asset pricing regression approach. The starting point is that the spread is determined by expected default risk and credit history. We also add the BAA rated US corporate bond spread to capture the risk appetite of the banking sector. As default risk is an unobserved variable, we replace the expectation term with its realization and a prediction error. This creates an identification problem, as the realization is correlated with the prediction error. Assuming that expectations are rational, we can use any variable (debtor characteristics) available at the time of pricing as a valid instrument. Identification is thus based on the following intuition: with the exception of credit history (dynamic contracts) and the BAA spread (risk appetite), all fundamentals affect the spread only through their impact on the probability of default. Notice that the reduced-form regression of such a framework simply regresses the spread on various debtor characteristics.

In the reduced-form estimation, we find that both recent and distant default history have a significant positive influence on the spread, but the inclusion of country fixed effects is necessary. The conclusion of our benchmark structural-form specification is that future default risk, an overall risk appetite measure (the BAA-rated US corporate bond spread) and recent default history can robustly and meaningfully describe the spread.

In the structural form, the coefficient estimate for default risk is around 0.3. This means that if we increase our default risk measure from its median to the 90th percentile, the spread goes up by 19 basis points. Compared to the average spread of 133 basis points, it can be considered as sizable. For credit history, the structural-form estimate is very similar to the reduced form. Its magnitude is around 7.1. For this variable to affect the spread in a quantitatively meaningful way, we need to look at countries with an extremely poor record: if we increase the credit history indicator from its median to the 95th percentile, the spread goes up by 15 basis points, but for certain countries in our sample, there is a 31-67 basis point direct punishment component in the spread (keeping default risk unchanged). In a simple calculation, we show that the implied punishment for arrears is large enough to ensure timely payment in case a country has at least a 0.89 subjective discount factor (assuming a 0.95 world discount factor). Overall, the structural-form estimation offers strong statistical and economic evidence that credit history has a dominantly direct effect on loan spreads. Though we discuss some alternative explanations, we argue that the most plausible interpretation of this result points to the presence of dynamic contracts in sovereign bank lending.

The structure of the paper is the following: the second section comprises a review of the theoretical and empirical literature on the role of credit history. The next part develops the empirical strategy of the structural-form asset-pricing setup. The description of data, variables and the main econometric problems is presented in the fourth section. The fifth part describes the reduced- and structural-form results, while the last section concludes.

2 Related theoretical and empirical literature

The sovereign risk literature has identified many channels through which a default can influence future borrowing terms. One interpretation is based on signaling: a default reveals some information about the debtor, which hurts its future outcomes (borrowing terms or third party decisions like private investment), so the debtor avoids default in order to send a favorable signal about its fundamentals (as in Sandleris, 2008) or type (as in Eaton, 1996).

Alternatively, one can view sovereign borrowing as a dynamic contract between borrowers and lenders. In this case, there is an implicit or explicit agreement on lending and repayment terms, and any deviation (default) would initiate some punishment. Reputation as a repayment incentive then refers to the case of indirect punishments, like exclusion from future borrowing. In most cases (like Eaton and Gersovitz, 1981, Kletzer and Wright, 2000, Wright, 2002, Yue, 2009), this punishment is an out-of-equilibrium threat. In some models this is already sufficient to ensure full repayment (Eaton and Gersovitz, 1981, Kletzer and Wright, 2000, Wright, 2002). Consequently, there is no default or punishment in equilibrium. In Yue (2009), the punishments are the threat points of the renegotiation process, so there is default in equilibrium, but no punishment. Note that being an out-of-equilibrium threat need not imply the lack of credibility: both in Kletzer and Wright (2000) and Wright (2002), the threats are subgame perfect and renegotiation proof.

There are some models with default and/or punishment along the equilibrium path. In most cases, however, default is a consequence of incomplete markets (debt contracts are non-contingent). Examples include Yue (2009), Sandleris (2008), and Arellano (2008). The recent literature on unsecured household debt (for example, Chatterjee et al, 2007) also allows for default, with an exogenously set bankruptcy procedure (unrelated to reputation) as the punishment. There is equilibrium default in the investment model of Hopenhayn and Werning (2008), but there is no direct punishment after a default, the relationship simply ends.

Relaxing the assumption of perfect observability of parties' actions (like bank fees kept confidential), the theory of repeated games with imperfect monitoring implies that even contingent contracts could feature occasional episodes of punishment and potentially "default" as well: in the cartel literature, Green and Porter (1984), and Tedeschi (1994) establish theoretically that collusive and non-cooperative behavior is mixed, even in the absence of actual cheating (default) in the cartel.[1]

In Kovrijnyikh and Szentes (2007), there is no explicit breaking of the contract. Instead, they have a "debt overhang" situation, when the borrower cannot fully repay in a given period, which gives a de facto monopoly power to the incumbent lender. After "writing off" part or all of the initial debt, the incumbent can neglect other lenders and can extract some monopoly rents ("endogenous exclusion from competitive markets"). If one interprets a debt overhang as a full write-off of the initial debt (which is a sunk cost for the incumbent), then the monopoly contract in fact implies an interest rate giving extra profits to the incumbent.

The one-period loan contract interpretation of Kletzer and Wright (2000) also features nonzero lender surplus after a default. Each period starts with some nonnegative repayment R, followed by a zero expected profit loan contract of size L. The repayment is the lender's surplus in the general formulation. Suppose that there is default in a certain period, so R is not paid and L is not granted. Next period, the lender expects the borrower to give her the highest (state contingent) R in order to continue the relationship. One can interpret this punishment as a new loan of $L - R$, but for an expected present discounted value of repayment equal to $L > L - R$.

It is not obvious that a competitive capital market is compatible with nonzero lender surpluses after a default, as a competing lender might be tempted to offer a cheaper loan. In dynamic relations, however, positive surpluses can be maintained by repeated lender interactions. In Kletzer and Wright (2000), the punishment is compatible with external competition, due to a "cheat the cheater" response of the other lenders. This leads to

[1] Abreu, Pearce and Stacchetti (1990) generalize these results to discounted repeated games.

an implicit seniority of preexisting loans.[2] Kovrijnyikh and Szentes (2007) assume such a seniority explicitly, which gives the incumbent lender monopoly power after a debt overhang. The common feature is that preexisting debt limits the impact of outside competition, and allows for punishments which give positive surplus to the lender even in the case of potential entrants. An alternative mechanism is described by Wright (2002): for syndicated loans, in which each bank has a share of the profits, then each bank's incentive to maintain a good reputation in this cooperation makes them tacitly collude in punishing a country in default.[3]

Switching to empirics, there is a diverse literature aimed at detecting behavior in line with dynamic incentives. In the economic history literature, Greif (1993) looks at the coalition of Maghribi traders in medieval times, and finds that their merchant-agent relationships evolved consistently with a repeated game with imperfect monitoring. Milgrom, North and Weingast (1990) also point to the role of self-enforcing contracts in medieval trade. Porter (1983) finds empirical support for repeated games with imperfect monitoring (switches between collusive and noncollusive behavior in the US railroad industry), while Levenstein (1997) performs a similar analysis for the bromine industry in the US.

Focusing on sovereign borrowing, there is some direct evidence on the repayment incentives of a sovereign debtor. This literature is surveyed and discussed in Panizza et al (2009). The more relevant issue for our discussion is how a country's credit history affects the borrowing cost of a sovereign. Eichengreen and Mody (1999) use data on 4500 loans over the 1991-1997 period and employ a pooled OLS regression, corrected for sample selectivity. They notice that a history of debt reschedulings has a weak positive effect on the probability of an issue while it significantly increases the spread that successful issuers pay.

Ozler (1993) is an important contribution to the issue of reputation, which has been cited by many theoretical papers as the main evidence of an effect of repayment history on credit terms.[4] She uses data on 64 countries for the period 1968-1981, which was one of rapid international lending expansion. The econometric technique is a pooled OLS regression with time-specific dummies. Her main results are that the spread is influenced by relatively recent repayment history (which she identifies as the 1930s through 1960s), but not by distant history (before 1930).

Reinhart et al (2003) provides a documentation of the effect of repayment behavior on sovereign debt. Employing a cross-sectional regression with multiyear averages of measures

[2]Drelichman and Voth (2011) use archival data to analyze sovereign lending to King Philip II of Spain, 1556-1598, a period characterized by the borrower's repeated defaults. They find evidence in favor of lender's coalitions employing punishment strategies that were sustaining lending based on the "cheat-the-cheater" mechanism described in Kletzer and Wright (2000).

[3]Coordination is also an issue in Kovrijnyikh and Szentes (2007): if each lender represents a group of banks, than the group must form a credible syndicate when exercising its monopoly power.

[4]See for example Eaton (1996) or Obstfeld and Rogoff (1996, p. 379).

for default risk, history of repayment, inflation rates and external debt as controls, they find that a history of defaults weakens a country's ability to borrow large amounts on reasonable terms, because a bad credit history is reflected in lower creditworthiness (proxied by the country's credit rating).

Our analysis extends that of Ozler (1993) and Reinhart et al (2003) in two major ways. The first one concerns the treatment of country fixed effects. By including time invariant variables like dummies for repayment problems, Ozler can no longer have country fixed effects. We resolve this issue by constructing a continuous measure of recent default, which allows us to use both country fixed effects and default history indicators.

The second, more important contribution is a structural and causal empirical approach. The ability to distinguish between different channels of influence is particularly important for the credit history case. As we argued before, there are two ways in which history could affect the spread. Looking only at the reduced-form results – as Ozler (1993) and Reinhart et al (2003) do – one cannot separate the two effects.

3 The empirical strategy: identification in a structural-form asset pricing regression

The starting point is that the spread reflects perceived risks and potential extra effects (like a punishment surcharge for past defaults, or the overall risk appetite of the banking sector). Such a specification thus involves latent expectations of the risk(s) based on information at the time of pricing. Three main solutions have been adopted to overcome this issue. One widely used approach has been to assume specific functional form relations between the spread, the risk probability and the economic fundamentals to get, by substituting one into the other, an estimable reduced-form equation. Examples include Edwards (1986), Ozler (1993), Eichengreen and Mody (1999), Easton and Rockerbie (1999). Another solution has been to use proxies for the probabilities, like credit ratings (Kamin and von Kleist, 1999). A third approach has been to use multiple issues of the same borrower, assuming a common default probability (Cumby and Pastine, 2001). All these methods suffer from a common problem: they cannot identify more than one source of risk and test for a systematic extra effect of certain country characteristics.

We adopt a rational expectations approach and employ the errors in variables method (EVM) as a solution for these problems. Start from the representation of the structural form of the pricing equation:

$$s_{it} = \alpha + \beta R_t + \lambda E(d_{it}|Z_{it,}R_t) + \Theta X_{it} + \varepsilon_{1it}, \tag{1}$$

where s_{it} is the spread paid by country i on loans obtained at time t, R_t is the benchmark interest rate, d_{it} measures the risk of default on the loans obtained by country i at time t, X_{it} is a vector of various extra factors, and Z_{it} contains information available at the time of pricing. Notice that Z_{it} contains all of the X_{it} variables as well. The error term ε_{1it} is orthogonal to any time t information (Z_{it} and R_t).

The linearity of (1) can be derived from risk-neutrality,[5] profit maximization, and assuming partial default (with probability p) on the principal (denoted by x) but not on the interest (r):

$$(1-p)(1+r)+p(x+r)=1+R,$$

which implies $s=r-R=p(1-x)$. Based on this, the coefficient λ of default risk in (1) would reflect potential losses from default (expected haircuts).[6]

According to the EVM method, one replaces the expectation term in (1) with its realization:

$$d_{it}=E(d_{it}|Z_{it,}R_t)+\varepsilon_{2it}. \tag{2}$$

Given rational expectations, $E(\varepsilon_{2it}|Z_{it,}R_t)=0$, and thus equation (1) becomes:

$$s_{it}=\alpha+\beta R_t+\lambda d_{it}+\Theta X_{it}-\lambda\varepsilon_{2it}+\varepsilon_{1it}. \tag{3}$$

Now, d_{it} is not orthogonal to the compound error term, since it is not orthogonal to the prediction error ε_{2it} (see equation (2)) and ε_{1it} (possible simultaneity problem[7]). But according to the EVM approach, one can use the information set $\{Z_{it},\ R_t\}$ as valid instruments, since this set is correlated with the default event (from the prediction equation (2)) and uncorrelated with the error term (from the rational expectations assumption and the pricing equation (1)).

[5] Risk-neutrality is clearly a strong assumption, which we are relaxing in two ways. First, we allow for an additive global risk-aversion term by including the US BAA spread on the right hand side (motivated by the findings of Longstaff et al, 2011). Second, as long as the model-implied risk premium is proportional to the probability of default (like in Arellano and Ramanarayanan, 2008, and Borri and Verdelhan, 2010), then the coefficient of the probability of default combines the effect going through expected default and through the risk premium. See Section 5.4 for further discussion.

[6] The usual approach in the empirical literature is to assume that the bond either pays the benchmark rate plus the spread, or zero. In that case, risk-neutrality implies that the log of the spread is a (unit coefficient) linear function of the log odds ratio $\frac{p}{1-p}$. This is not suitable for our purposes: as explained by equation (2), we want to replace p with its realization d, but that is a zero-one variable, for which the log odds ratio is undefined. Moreover, actual defaults are more complicated than a delivery of zero payoff for a one period debt instrument, so the usual approach also involves strong simplifications. Our default assumption only serves to motivate the linearity of (1).

[7] This means the following. A pricing error ε_{1it} can lead to a higher future default loss d_{it}, the conditional expectation of which is a right hand side variable in the pricing equation. As the pricing error is assumed to be orthogonal to Z_{it} and R_t, it means that those variables remain valid instruments for this source of nonorthogonality as well.

In particular, our benchmark specification is the following

$$s_{it} = \alpha + \beta R_t + \lambda d_{it} + \theta_1 recentdef_{it} + \theta_2 BAAspread_{it} + \varepsilon'_{1it}, \tag{4}$$

where $recentdef_{it}$ is our measure of credit history, and $BAAspread_{it} = BAAspread_t$ is the BAA rated US corporate bond spread, capturing the risk appetite of the banking sector.

Due to the properties of the instrumental variables estimation, this method provides consistent though not fully efficient estimates even when the information set is incomplete or the functional form of the prediction equation is unknown (this was already emphasized by Wickens, 1982). In particular, estimates of the structural form (1) are robust to potentially omitted fundamentals from the set Z_{it}, while the reduced form (5) is subject to omitted variable bias. This is a major advantage, given the potential sensitivity of empirical results to functional form assumptions, selectivity bias and omitted variables. The key element is whether the fundamentals are sufficiently correlated with the default variable. If they are, then they can be used as valid instruments in the pricing equation, without having to specify the default prediction equation.

It is important to stress that a variable included in X_{it} affects the spread through two channels: through an impact on predicted future default $E(d_{it}|Z_{it,}R_t)$, and a direct effect through Θ. The total effect is captured by the reduced-form equation

$$s_{it} = \alpha' + \beta' R_t + \Gamma' Y_{it} + \Theta' X_{it} + \varepsilon_{it}, \tag{5}$$

where $Y = Z \backslash X$. Denoting the linear conditional expectation of d_{it} by

$$d_{it} = \alpha'' + \beta'' R_t + \Gamma'' Y_{it} + \Theta'' X_{it} + \varepsilon_{2it},$$

the structural form (1) imposes the following restrictions on the reduced form (5):

$$\alpha' = \alpha + \lambda\alpha'';\ \beta' = \beta + \lambda\beta'';\ \Gamma' = \lambda\Gamma'';\ \Theta' = \Theta + \lambda\Theta''. \tag{6}$$

This immediately shows the decomposition of the total effect Θ' into the direct effect Θ and the indirect effect $\lambda\Theta''$. Moreover, it also illuminates the way we identify this decomposition: Γ'' and Θ'' are obtained from the prediction equation for d_{it}, the risk parameter λ is identified through the restriction $\Gamma' = \lambda\Gamma''$, while Θ is obtained as $\Theta = \Theta' - \lambda\Theta''$. This is exactly what an instrumental variables (IV) estimation of equation (1) does in one step. Intuitively, our identification is based on the following. With the exception of credit history (dynamic contracts) and the BAA spread (risk appetite), all fundamentals affect the spread only through their impact on the probability of default. They can be thus utilized to separate the direct impact of credit history $\theta_1 recentdef_{it}$ from its indirect impact through d_{it}.

4 Data, variables and estimation issues

The choice of the time period was mostly driven by Ozler's (1993) observation that a period of market expansion is needed to distinguish the impact of an individual borrower's repayment history from the impact of a widespread panic. Thus, we use the period 1973-1981, which witnessed particularly intense syndicated bank lending to sovereign borrowers. In fact, bank loans were the dominant source of sovereign capital flows in the 70s, which was no longer true after the Debt Crisis. The initial dataset contains the spread (over the 1-year LIBOR) on 757 commercial bank loan contracts denominated in dollars, to 46 developing countries and were obtained from various issues of the World Bank's "Borrowing in International Capital Markets".

As we have no access to contract-level characteristics of loans or their future repayment patterns, we average over all contracts of the same country at a given time period. Since the economic fundamentals are mostly available at the annually frequency, we construct yearly measures for the spread. Just like Easton and Rockerbie (1999), we use a weighted average of the original spreads, using as weights the loan quantities and maturities. As an alternative, we also discuss results which use an average spread weighted by loan quantities only (like Edwards, 1986 and Ozler, 1993). Using maturity in the weighting allows for taking into account that the spread of a longer maturity debt influences average credit terms to a larger amount than the spread of a shorter maturity loan.[8] This transformation means that we are

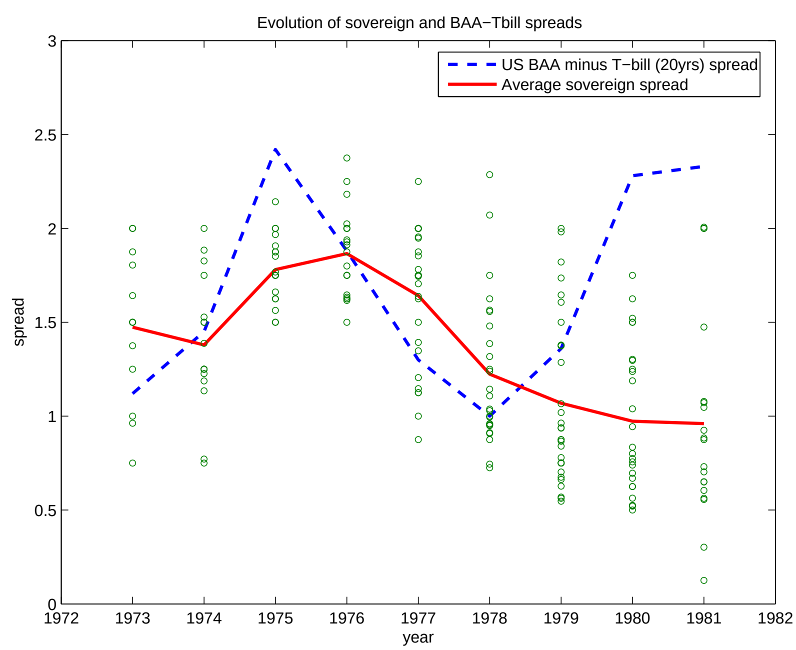
[8] The World Bank's Global Development Finance database also reports average interest rates weighted by quantities. From a theoretical point of view, however, it is not obvious what the correct weighting should be. Suppose that a loan contract specifies an interest payment stream of a constant spread over the benchmark yield $r_t + s$ for a period T, at the end of which the full principal q is repaid (notice that actual contracts might specify a different schedule for principal repayments). It is straightforward to see that the present discount value (PDV) of such a loan contract is $qs[1 - \exp(-\int_0^T r_\tau d\tau)] = qs(1 - R_T^{-1})$. Here r_τ is the instantaneous required interest rate, while R_T is the required yield from time zero to T. Having two loan contracts (q_1, s_1, T_1) and (q_2, s_2, T_2), the equivalent contract ($q_1 + q_2, s', T'$) then must satisfy $q_1 s_1(1 - R_{T_1}^{-1}) + q_2 s_2(1 - R_{T_2}^{-1}) = (q_1 + q_2)s'(1 - R_{T'}^{-1})$. This defines s' for a given T'. In order for this to be a weighted average of s_1 and s_2, T' must be such that $R_{T'}^{-1}$ is the quantity weighted average of $R_{T_1}^{-1}$ and $R_{T_2}^{-1}$. Then s' is a weighted average of s_1 and s_2, with relative weights $q_1(1 - R_{T_1}^{-1})$ and $q_2(1 - R_{T_2}^{-1})$. For $T_1 \neq T_2$, this in general requires a detailed knowledge of the entire benchmark yield curve. If the yield curve is flat ($r_t = r$), R_{T_1} and R_{T_2} are not too large (in the sense that $R_{T_1}^{-1}$ and $R_{T_2}^{-1}$ are well approximated by $1 - R_{T_1}$ and $1 - R_{T_2}$), then the spread on the equivalent contract is precisely the quantity and maturity weighted average of s_1 and s_2.

left with 201 yearly observations.

Data availability (arrears, country fundamentals) and the need of first differencing further reduce the working sample, to 158 observations from 36 countries. Finally, we excluded the most extreme values of our constructed default variables.[9] The final sample for which we report the results is 151 observations and 35 countries.

Figure 1 illustrates the evolution of the sovereign spreads in the sample, together with the average BAA-rated US corporate bond spread (taken as the difference between the BAA corporate bond yield and the fixed 20-year maturity Treasury yield series from the Federal Reserves' website). The spread is measured in percentage points. There are several aspects of interest in the visual comparison of the two spreads. First, the figure shows that sovereigns pay similar spreads to BAA-rated US companies. Second, the variation in spreads is very large in 1981, suggesting that commercial banks were distinguishing between the borrowers, even before the "unexpected" debt crisis of 1982. Third, visually, one can notice that the sovereign spreads and the BAA spread strongly comove until 1978, and then move in the opposite direction between 1979-1981.

Figure 1: The evolution of sovereign spreads for 1973-1981 (compared to the average BAA-rated US corporate bond spread; in percentage points)



[9]Specifically, we eliminated 3 observations for the recent default that were more than 10 times the sample standard deviation and the bottom and top 1% for the future default variable.

4.1 Repayment history indicators

There is no clear indication from theory regarding the choice of the repayment history variable. Moreover, in the context of repeated games with imperfect monitoring, punishment is invoked by some imperfect indicator of cheating, and not an outright default episode. In general, the repayment history variable should still be related to the overall loss creditors incurred due to repayment problems. Recognizing that any indicator is merely a proxy, our choice is restricted by data availability. In particular, binary indicators of repayment problems are available both historically and recently; arrears data is reported by the World Bank from 1971; while the size of debt forgiveness and reschedulings are reported by the World Bank only from 1989. For this reason, we use binary indicators for capturing "distant" repayment history and arrears data for "recent" history.

There are reasons to believe that recent and distant history have a different effect. Indeed, Ozler (1993) finds that repayment difficulties happening before the 1930s did not significantly matter for spreads in the 1970s, while those happening afterwards did. In the models of Yue (2009), Kovrijnykh and Szentes (2007), and Benjamin and Wright (2009), it is also recent default (arrears) that matters; in fact, once a country eliminates its arrears, it gets a clear credit history. From the point of view of dynamic contracts, we expect that it is mostly recent history that matters for potential punishments.

For distant history we use an indicator of the presence of default or rescheduling of bank loan debt to official creditors in the period 1940-1970. This dummy variable was constructed based on Ozler (1993), which includes data for 1956-1968, and Lindert and Morton (1989), which refers to the period 1940-1970. Table 1 presents its summary statistics, the number of observations and countries that had repayment problems. It is important to note that the indicator has significant variation to be able to identify the effects. The mean for this variable is high, and shows that around 35% of the countries in the sample had some repayment problems during 1940-1970.

Table 1: The "distant" default variable[a]

Variable: distant default dummy	Total observations /countries	Obs. with 1	Mean	Std. Dev.
Whole sample	201/37	70	0.348	0.477
Restricted sample	151/35	56	0.371	0.484

a Constructed as a dummy variable for repayment problems on loans for 1940-1970. The dummy takes the value 1 for a repayment problem.

While this indicator is very similar to that used in Ozler (1993), the indicators reflecting recent history are our own. Their construction allows including a continuous variable instead of a dummy to reflect past repayment behavior, enabling to control for country fixed effects and still include a default history measure. We construct this indicator from data on private arrears (both interest and principal) on long-term debt outstanding, available since 1971 from the Global Development Finance CD-ROM. As Cline (1984) notes, debt reschedulings are usually preceded by the accumulation of arrears, thus their presence and size can be a good indicator of potential creditor losses.

Our measure of recent default is based on the sovereign borrower's stock of accumulated arrears. The benchmark time t variable controlling recent default is the stock of arrears at time $t-1$ divided by the loan amount at time $t-1$. There are several reasons to favor this measure. One, it uses the most recent information available to the investor at the moment of pricing on the amount of arrears accumulated. Second, by normalizing the stock of arrears with the amount disbursed, we are consistent in constructing the recent and future repayment problem indicators as proportional to the loan amounts. Third, as we will discuss in a later section, this measure allows for a meaningful economic interpretation of the estimated coefficient on the recent default indicator as being related to the country' subjective discount factor.

4.2 Future default variables

To measure how future default risk is priced into the sovereign spread, the future default variable should closely reflect the realization of proportional losses on a loan. As demonstrated by Sturzenegger and Zettelmeyer (2007), and Benjamin and Wright (2009), precise measures of realized repayments are very hard to compute for sovereign debt. Easton and Rockerbie (1999) argue that arrears are more indicative for repayment problems than default or rescheduling indicators. Based on these, we construct our future default measure by using again GDF data on arrears: from any given period onwards, we cumulate private arrears for a period similar to the average maturity (the detailed procedure is explained in the Appendix). To reflect relative losses, this sum needs to be normalized by the total amount of loans outstanding, corresponding to the given period.

Table 2 provides some brief descriptive statistics of our benchmark choice for future and recent default. For recent default around 75% of the observations are equal to 0 for both the full and the reduced sample; this number is around 20% for the future indicator. The difference is due to more frequent arrears after 1981, but it shows that there were still countries that were not accumulating arrears in this period.

Table 2: The "recent" and "future" default variables[a]

Variable Variable	Total obs. /countries	Obs. with 0	Mean	Std. Dev.	10%	Median	90%
Recent default[b]	198/37	152	0.0037	0.015	0	0	0.0047
Restricted sample	151/35	112	0.003	0.011	0	0	0.0031
Future default[c]	201/37	43	0.143	0.357	-0.0001	0.0015	0.465
Restricted sample	151/35	29	0.149	0.340	-0.0005	0.002	0.659

a Constructed as continuous variables based on arrears data. A zero means no repayment problem.
b The indicator uses the time t-1 stock of private arrears and divides it by the time t-1 loan amount. Information refers to the whole sample.
c The indicator adds private arrears for 8 years in the future and divides them by the loan amount. Information refers to the whole sample.

4.3 Economic fundamentals

The first set of fundamentals consists of variables that we expect to influence the spread only indirectly. In other words, they can be significant in the reduced form and/or the prediction equation, but they should influence the spread only because they influence (predict) future default. The sources for these variables are Global Development Finance, International Financial Statistics and World Development Indicators. Besides data availability, we are following most of the literature in considering a wide set of country-specific economic fundamentals that could potentially affect the pricing of sovereign loans.[10] Out of the set of potential economic fundamentals we retain for the benchmark specifications the variables that are influential in their reduced-form effect on the spread.

We construct two additional variables that are related to the international financial environment of a country. The first, "Repeated borrowings" is designed to capture the importance of relationship banking.[11] It is constructed by cumulating the number of years in which the borrower received loans. The second, "Proportion of countries with arrears in the region" is aimed at capturing a regional contagion effect from one country going into arrears.[12] It is obtained by dividing the number of countries with arrears from the same

[10] We consider the set of following variables as candidates: debt to GDP ratio, reserves to imports ratio, debt service to exports ratio, current account per GDP, exports to GDP ratio, savings to GDP ratio, growth of per capita GDP, growth of gross investment, GDP per capita, inflation, terms of trade changes, credit to private sector per GDP.

[11] Ozler (1992) finds that such a measure is significant in explaining variation in sovereign loan spreads for developing countries over the 1968-1981 period.

[12] See Edwards (2000) for an overview of contagion.

region by the total number of countries in that region.[13]

The second group of fundamentals contains variables that can potentially have a direct effect on the spread, not only through future default. It is quite important to control for such factors, otherwise a direct effect of recent default might be picking up risk appetite, for example. The first candidate is the benchmark interest rate, for which we use the LIBOR USD 1-year rate. Besides recent default and the benchmark rate, there are two major factors that can easily have a direct impact on sovereign debt prices. One is the overall risk appetite (or risk aversion) of the market,[14] while the other is market liquidity.[15] In case of new sovereign *loan* disbursements from the 70s, liquidity should not be an issue: these are primary issues of loans, without active secondary market trading. To control for risk appetite, we use the difference between the US BAA corporate bond yield and the fixed 20-year maturity Treasury yield.

4.4 Estimation issues

Both the reduced- and the structural-form specification are subject to two major econometric problems: the need to control for country fixed effects, and the validity of the strict exogeneity assumption.[16] The Appendix contains an in-depth discussion of how we handle these issues, here we just briefly outline our strategy.

In the reduced form, we use a pooled OLS, a random effects, a fixed effects and first differencing estimators. In the structural form, we use the first difference estimator with appropriately instrumenting the future default variable: it eliminates the individual effects, and the right choice of instruments resolves the endogeneity problem caused by the prediction error in a way that requires only the sequential exogeneity assumption. The appropriate instruments include the first and/or second lags of the regular instruments (time t information). Using as instruments the levels of the variables (as opposed to the second difference) leads to more precise estimates, but at the cost of making the direct comparison with the reduced-form results of Table 3 more difficult.

The whole structural-form estimation framework is based on the validity of the instruments: they should be correlated with the instrumented variable and also uncorrelated with the error terms. For this reason, we report various measures that summarize the

[13] We also construct a dummy variable that reflects whether the country has gained sovereignty before or after 1930. The source for this information is Ozler (1993) where she finds a significant effect of this variable in pricing sovereign bank loans.

[14] As shown for sovereign CDS spreads by Longstaff et al (2011).

[15] As shown for US corporate bonds in Elton et al (2001).

[16] Strict exogeneity means that the idiosyncratic error terms, conditional on the individual effect, are uncorrelated with past, present and future values of the regressors.

first-stage regression, and an overidentification test that is appropriate in a setting with heteroskedasticity and autocorrelation (the Hansen's J statistic), and it is consistent even with intra-cluster correlation.

5 Results

5.1 The reduced form

The results of the reduced-form estimation are presented in Table 3 and refer to four specifications. The left hand side variable is a quantity and maturity weighted annual spread. The recent default variable used is the stock of arrears at t-1, divided by the time t-1 loan amount. As Column (1) shows, most of the explanatory variables are significant and have the expected sign in the pooled OLS specification. In particular, the coefficient of both distant and recent default is positive and significant . The significance of the distant default replicates the main finding of Ozler (1993). When running panel data specific regressions like fixed effects (Column 2) and random effects (Column 3), the results do not change substantially except for two variables, private sector credit to GDP and proportion of countries in the region with arrears. Column (4) is a specification in which the variables are first-differenced. The estimates' precision is smaller, due to the fact that there is less time-variation and that by first differencing there is one observation not used for each country. Still, the effect of recent default is positive and (marginally) significant, even after controlling for country effects.

In conclusion, we find that recent default and many additional economic fundamentals significantly influence the spread, even after controlling for country specific effects. There is also some indication that the significance of distant default might be a consequence of omitted country effects.[17] This latter finding questions by itself the existing empirical literature (such as Ozler, 1993 and Reinhart et al, 2003), on the role of distant default history in pricing sovereign loans.

5.2 The structural form

The results of our benchmark specification (4) are presented in Table 4. As motivated in Section 4, we choose as our main structural-form estimation the one in first differences, since

[17]Though fixed effects are not compatible with our distant default variable, the comparison of the distribution of country effects across groups with different default history might convey some information. Indeed, the results from the reduced-form estimations suggest that defaulters (categorized through the distant default dummy) are not being charged significantly higher spreads. Specifically, the difference between the means of the two groups is not significantly different from zero.

Table 3: Reduced-form estimation: the determinants of the spread[a]

	Pooled OLS (1)	Fixed Effects (2)	Random Effects (3)	First Difference (4)
Benchmark yield	−0.11	−0.11	−0.11	−0.099
	(−11.33)**	(−11.79)**	(−13.04)**	(−11.48)**
US BAA-TB spread	0.23	0.24	0.24	0.24
	(7.46)**	(6.28)**	(6.49)**	(8.79)**
Recent default	8.33	7.89	7.13	6.45
	(3.16)**	(2.81)**	(2.93)**	(1.35)
Distant default	0.11	-	0.07	-
	(1.78)**	-	(0.7)	-
Reserves to imports	−0.92	−0.56	−0.71	−0.38
	(−4.53)**	(−3.32)**	(−4.75)**	(−3.43)**
Private credit to GDP	−0.006	−0.003	−0.006	0.003
	(−2.58)**	(−0.87)	(−2.12)**	(0.46)
Debt to GDP	0.26	0.29	0.27	0.27
	(1.68)*	(1.54)*	(1.78)**	(1.13)
GDP growth	−0.52	−0.77	−0.72	−0.38
	(−1.88)**	(−2.58)**	(−2.51)**	(−1.77)**
Countries with arrears	0.42	0.53	0.52	−0.26
in the region	(1.95)**	(1.25)	(2.34)**	(−0.7)
Repeated borrowings	−0.016	−0.028	−0.026	−0.022
	(−0.8)	(−1.87)**	(−1.96)**	(−1.01)
New sovereign	0.046	-	0.044	-
	(0.66)	-	(0.48)	-
Constant	2.78	2.92	2.9	–
	(9.87)	(7.96)	(8.58)	–
No. of obs	192	192	192	155
R^2	0.66	0.64	0.65	0.55

a The t statistics are in parentheses; the standard errors are corrected for clustering at country level. *, ** denote 0.2 and 0.1 significance levels, respectively.

it provides the framework that allows us to make correct inference on the overidentification test and the right hand side variables. We also briefly comment on some results from the level specification, as they can shed light on the channel decomposition of distant default. Overall, there are three important findings we discuss here: the influence of the future default indicator, the coefficient of the benchmark yield, and most importantly, the effect of the recent default indicator.[18]

Starting with the first, the future default's point estimate is around 0.3, significant at the 10% level. Although the mean of this indicator is just 0.15, this is not very indicative of its influence, because the variance is large and for many countries the indicator's value is around 0.5 and even 1. If we consider an increase in the indicator from its median to the 90th percentile, then this would raise the spread by approximately 19 basis points. Consequently, the coefficient can be considered as sizable, as the sample mean of the spread is 133 basis points. This is an important finding, because it suggests that expected default risk was priced in the lending decision and that the debt crisis of 1982 was, to this extent, "anticipated". Nevertheless, it remains true that for many countries this prediction was correct mainly in sign, and less in the size, as the quantitative effect of the future default risk is small.

A second general feature is the significance of the benchmark rate, which was also present in the reduced-form results. Given that the spread is defined as the difference between the loan rate and the LIBOR rate, this result is equivalent to the finding that the loan rate responds less than one-in-one to the world interest rate (the reaction coefficient is around 0.9). This conclusion is found also in Eichengreen and Mody (1999) and Uribe and Yue (2006).

The central results concern the channels of influence of the economic fundamentals, particularly that of the recent default indicator. Our benchmark result reported in Table 4 suggests that the recent default indicator, included as an extra right hand side variable in the pricing regression, has a significant positive direct effect on the spread, above the indirect effect captured by the instrumenting of future default with information available at the time of pricing.

The estimate on the recent default measure is not very sizeable. Still, our analysis shows that this effect is statistically significant and in some cases economically too. For this variable to affect the spread in a quantitatively meaningful way, we need to look at countries with an extremely poor record: if we increase the credit history indicator from its median to the 95th percentile, the spread goes up by 15 basis points, and for certain countries in our sample, there is a 31-67 basis point direct punishment component in the spread.

[18]One can also notice that the BAA spread has a significantly positive coefficient, indicating that an increase in the overall risk appetite of the (US) financial sector pushes up sovereign loan spreads as well.

Table 4: Structural-form estimation: the determinants of the spread[a]

	Estimation method[b]
Future default	0.291
	(1.83)**
Benchmark yield	−0.109
	(−11.91)**
US BAA spread	0.233
	(7.97)**
Recent default	7.22
	(1.75)**
First stage relevance:[c]	
Partial R^2 for future default	0.254
Kleibergen-Paap rk Wald stat p-val[d]	0.006
Structural form:	
Overidentification test p-value[e]	0.742
Number of observations	151

a The t statistics are in parentheses; the standard errors are corrected for clustering at country level. *, ** denote 0.2 and 0.1 significance levels, respectively.

b The dependent and explanatory variables are first differenced. The future default variable is instrumented by the first lag of the following variables: benchmark yield, debt/GDP, reserves to imports, GDP growth, proportion of countries with arrears in the region, repeated borrowings, private sector credit/GDP, new sovereign dummy and a distant default dummy.

c The reduced-form regression of the instrumented indicator on the full set of instruments.

d The Kleibergen-Paap rk Wald test of the null hypothesis that the equation is underidentified.

e The Hansen J-statistic.

It is important to note however, that this type of calculation refers to only a one period ahead increase in the spread caused by recent arrears. In fact, this punishment lasts potentially for more periods. More specifically, given that the recent default indicator reflects a stock of accumulated arrears, the effect lasts until this stock is eliminated. As Tedeschi (1994) suggests, even a punishment that is relatively small in each period can deter deviations (in his case: cheating in the cartel), as long as it can last for a potentially very long time. As an extreme case, if the punishment threat is sufficient to deter default completely, then one would not observe any punishment at all. Consequently, even a small price punishment is indicative for the presence of dynamic contracts.

Using a simple calculation we can show that the implied punishment for arrears is large enough to ensure timely payment in case a country has at least a 0.89 subjective discount factor (assuming a 0.95 world discount factor). To see this, we can analyze a simple tradeoff facing a country of whether to pay its debt now or pay it later. Recall that our benchmark recent default measure can be described as:

$$a_t = \frac{A_{t-1}}{Q_{t-1}},$$

where A is the stock of arrears and Q denotes loans outstanding. Let θ be the structural-form pricing coefficient on a_t. Now, suppose that at time t some payment P_t is due and that $A_{t-1} = 0$. If the country decides to postpone paying P_t until next period, then P_t becomes arrears for $t+1$:

$$a_{t+1} = \frac{P_t}{Q_t}.$$

In that case, the spread next period goes up by $\Delta s_{t+1} = \theta \frac{P_t}{Q_t}$ so the discounted payments at $t+1$ are:

$$\beta \left(\Delta s_{t+1} Q_{t+1} + \frac{P_t}{\bar{\beta}} \right) = \beta \left(\theta \frac{P_t}{Q_t} Q_{t+1} + \frac{P_t}{\bar{\beta}} \right),$$

where β is the country's subjective discount factor, and $\left(\bar{\beta}\right)^{-1}$ is the world benchmark interest rate. Letting $g \equiv Q_{t+1}/Q_t$, the country is indifferent between paying now or later if

$$\theta = \frac{1}{g} \left(\frac{1}{\beta} - \frac{1}{\bar{\beta}} \right).$$

Based on our results $\theta = 0.07$, so for a parametrization of $\bar{\beta} = 0.95$, $g = 1$, the lower bound on the country's subjective discount factor that ensures repayment is $\beta = 0.891$. The calculation shows that our estimated punishment effect can be interpreted as being approximately equal to the difference between the shadow interest rate and the world benchmark rate. From that perspective, an estimated value for this difference of 7% can be considered as economically meaningful and large.

Now we briefly discuss some diagnostic statistics for the structural-form pricing regression reported in Table 4. The null of homoskedasticity is not rejected by the Pagan-Hall general test statistic (a p-value of 0.9249), which suggests that, given its inferior small sample properties, an asymptotically efficient GMM is not needed. The Arellano-Bond test also rejects at 5% the null of serial correlation of the errors. The first-stage statistics indicate that the instruments are relevant, with a partial R^2 of 0.25. The Kleibergen-Paap rk Wald test of whether the excluded instruments are relevant (the null being that the equation is underidentified) rejects the null with a p-value around 0. The overidentification test of all instruments (in the form of Hansen J-statistic) fails to reject the null with a p-value of 0.74. These diagnostics suggest that the structural-form regression is well specified and the instruments are valid and relevant.

Next, we briefly touch upon the level specification. The only reason why this is relevant is that it can separate the effect of distant default by channels. Unfortunately, estimation in levels is plagued with several problems which render it unreliable. The presence of fixed effects calls for a FE or RE specification, where the strict exogeneity assumption fails, and it suffers from a very significant rejection of the overidentification test. The instrumented future default risk appears very insignificant and the point estimate is much lower than the one resulting from the correctly specified first differenced estimation.[19] We attribute these to the failure of the strict exogeneity assumption.

If instead we ignore fixed effects and run the pooled level IV, the results are broadly close to our benchmark first-difference specification: the instruments have a similar first-stage explanatory power, and there is a significant extra effect of recent default.[20] When the past default dummy is included as an extra right hand side variable, its coefficient is negative and not significant. Thus, this specification would suggest that the reduced-form effect of past distant default on the spread can be attributed entirely to future default risk, and not to a punishment channel. However, given the problems outlined earlier, we would not trust enough this level specification to make a strong conclusion.

5.3 Robustness

Here we discuss two alternatives in the construction of two key variables: the annual average spread and the future default variable. Starting with the first, we reran our estimations (the ones reported in tables 3 and 4) with a quantity-weighted average spread (instead of

[19] In the RE case, the point estimate on the future default indicator is closer, even though much less significant, to the first-difference estimation.

[20] One significant difference is that now the overidentification test rejects the null with a p-value of 0.09.

weighing by quantity and maturity). Reassuringly, the point estimates, their significance level, and the diagnostic statistics have remained almost identical. Regarding our future default indicator, we consider an alternative when we exclude a three year grace period from the summation in equation (7). This also leads to very similar estimates and conclusions.

5.4 Channel decomposition

An important exercise for our conclusion is to compare the point estimates for the recent default measure in the reduced- and structural-form estimation (recall equation (6)). We find that the estimated coefficient in the structural form is about 95% of that from the reduced form. Our identification strategy interprets this as evidence that the effect of recent default on the spread is almost entirely a direct one, pointing to a price punishment above increased future default risk. In other words, we find evidence for positive lender surpluses after a default episode, indicating the presence of dynamic contracts in sovereign bank lending.

We certainly acknowledge that the documented evidence of the positive lender surplus can be consistent with alternative explanations as well. Given the literature on asset pricing, and the fact that we analyze the pricing of new syndicated sovereign bank loans which rules out liquidity as an alternative, the main candidates are a risk premium and a systematic expectational error (overreaction). Here we discuss these two alternative explanations.

Turning to the risk premium possibility first, we argue along several lines why we prefer our maintained interpretation. First, Longstaff et al. (2011) argue that for sovereign credit default swap spreads, most of the variation in returns is driven by compensation for bearing global risk. We control for global risk appetite by including in our structural form the US corporate BAA spread. We find that this factor is positive and very significant in determining country-specific spreads, confirming the results of Longstaff et al. (2011). Our benchmark structural-form regression states that, controlling for this risk factor, our identified extra effect of past credit history is significant.

Second, as long as the model-implied (local) risk premium is proportional to expected default, then the estimated coefficient on the future default combines the effect going through expected default and through the risk premium. For example, in the sovereign bond pricing models of Arellano and Ramanarayanan (2008) and Borri and Verdelhan (2010), the risk premium is proportional to expected default.[21] If the risk premium is a nonlinear but monotonic function of expected default, our linear specification in the future default

[21] The latter find empirical evidence of risk premia in sovereign bond pricing. These premia are driven in their model by the positive correlation of the borrower's endowment with the lender's consumption growth. Under external habit formation, negative shocks to the borrower's endowment can then generate risk premia, while also predicting higher expected default probabilities.

probability is a local linear approximation of a nonlinear specification in the combined effect of future default and the proportional risk premium. As long as the linear approximation is acceptable, our identified extra effect of credit history is still uncorrelated with the risk premium. The reason is that, based on our identification, the extra effect is orthogonal to the indirect effect of credit history going through expected default and thus through the risk premium.

Besides risk neutrality, our other fundamental identifying assumption is rational expectations. This raises the possibility that what we interpret as a direct effect of default history might be simply an expectational error. The behavioral finance literature indeed presents ample evidence of biased expectations of investors, and proposes a market sentiment component of the stochastic discount factor (see, for example, Shefrin, 2008, page 3). Our focus, however, is on new, large and individual loan deals offered by large banks, which limits the role of individual mistakes and market sentiment. A Knightian uncertainty premium can also lead to "expectational errors" when investors (banks) exhibit ambiguity aversion. In such a framework, rational but ambiguity-averse lenders over-react to bad news about a country's fundamentals (see for example Epstein and Schneider, 2008). For that explanation to hold it must be that lenders systematically over-react to the information of past repayment problems by over-predicting future default. It is hard to find direct evidence against the possibility of a systematic overprediction of default probabilities, since we do not have any type of survey data to test for the subjective expectations version of our pricing regression. The argument that our identified extra effect reflects such pessimistic beliefs relies on the idea that in our sample the realized default happened on average less than expected by lenders. However, given that our data concentrates mostly on the period *before* the sovereign debt crisis of the 1980s, often viewed as an unexpected default event, it is unlikely that our sample reflects subjective default expectations that are even larger than the average realized ones.[22]

To summarize, our main structural-form result is that past repayment problems have a direct influence on the spread on top of the indirect influence through expected default. Conditional on our identification strategy, we interpret that evidence as a positive lender surplus following a default episode, indicating the presence of dynamic contracts in sovereign bank lending. We acknowledge other alternative interpretations of the structural-form result.

[22]Two further issues should be noted here. First, a similar argument as with risk aversion can be invoked: if the pessimistic expectations are proportional to the ones implied by the rational expectations assumptions, and the linear approximation of the nonlinear model is acceptable, then our extra effect is orthogonal to the difference in beliefs and cannot be explained by systematic expectational errors. Second, at a more general level, a systematic expectational error can still act as a repayment incentive similarly to a price punishment. If a country accumulates arrears, it will lead the lenders to over-predict expected default and charge a higher spread than implied by the objective probability distribution. This surcharge increases the borrower's expected financing costs.

We believe that out of these alternatives, the dynamic contracts story is the most natural candidate and well developed theory for studying a market such as the one we look at, namely syndicated sovereign bank lending.

6 Concluding remarks

We extended the existing empirical literature on the role of credit history in sovereign bank lending along two dimensions. One is that we used a continuous measure of past default, enabling us to control for country fixed effects. Our other, more important contribution is the empirical strategy that allows for the distinction of multiple channels of influence. This strategy is a structural-form asset pricing rational-expectations regression in which the spread may be influenced by multiple risks and factors. Using the errors in the variables method, we replace the expectation term with its realization and instrument the latter with information available at the time of pricing. We also add default history as an extra right hand side variable, to check whether it influences the spread not only through expected default risk, but has an extra effect on it.

The reduced-form estimation provides evidence that, after controlling for fixed effects, borrower and regional characteristics, both recent and distant repayment history are significant. This makes the result similar to that obtained by Ozler (1993) and implies that although country effects do matter, credit history does play a role in determining sovereign spreads. The structural-form regression provides strong evidence of an extra effect of credit history (a punishment effect) in *prices*, above the one going through predicted default loss. The finding that credit history matters beyond predicting future default points to the presence of dynamic contracts in sovereign bank lending, where repayment incentives are incorporated into future borrowing terms. The major structural-form specification includes the benchmark LIBOR interest rate, expected default risk, the spread for BAA rated US corporate bonds, and the recent default indicator. All these variables are significant and robust to different specifications.

In terms of default costs, we believe that in reality there is a complex mix of trading and political sanctions, spillovers to other transactions and relationships, signaling and dynamic contracting considerations. Our main result is that there is evidence of this last effect: an extra surcharge in loan prices. This points to the presence of dynamic contracts in sovereign bank lending.

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Appendix

Appendix A: Construction of the future default indicator

Let us denote the yearly GDF data on arrears by $A_{i,s}$, with s ranging from 1970 to 1989. As those arrears may refer to contracts from multiple years, we need to allocate them among contracts. Our approach involves the following steps. First, we split $A_{i,s}$ among contracts from years $s-1, s-2, \ldots, s-8$ in proportion to new disbursements in the corresponding year:

$$Y_{is,s-j} = A_{i,s} ND_{i,s-j} / (ND_{i,s-1} + ... + ND_{i,s-8}), \tag{7}$$

$j = 1, 2, ..., 8$ (the average loan maturity in our sample). Here $Y_{is,s-j}$ is the time s arrears on a time $s-j$ contract and $ND_{i,s-j}$ is the size of the time $s-j$ loan for country i. Then we cumulate the arrear fragments $Y_{it+1,t}, Y_{it+2,t}, ..., Y_{it+8,t}$ into which a time t contract goes over its lifespan:

$$D_{it} = Y_{it+1,t} + Y_{it+2,t} + ... + Y_{it+8,t}. \tag{8}$$

Thus, in order to recover the amount of arrears affecting an active loan contract in year t, we assume that all the time t arrears affect all the loans that have not matured yet, and the size of the contract specific arrear is proportional to the size of the contract. We motivate this by two arguments: one is that there is no information available on which contracts these arrears correspond to; and second, the assumption that these flows can be attributed to several preceding loans is consistent with the cross-default clauses that these contracts included. According to such clauses once a country enters into default or any repayment problem that constitutes a break on the contract with one lender, this will be treated as default also by the other creditors. We also consider the effect of the grace period, thus excluding from this summation contracts that are more recent than the average grace period of 3 years.

Appendix B: Estimation issues

The reduced form in a panel framework is:

$$s_{it} = \alpha + \beta R_t + \Gamma Z_{it} + c_i + \varepsilon_{it},$$

where ε_{it} is the idiosyncratic error term, Z_{it} are the economic fundamentals for country i known at time t and c_i is the unobservable individual effect.

The first major concern is that the usual pooled OLS estimates are incompatible with individual country effects. The usual procedure to correct for fixed effects is a fixed effects

(FE) or a random effects (RE) estimator. A key assumption behind both methods is strict exogeneity, which requires that the idiosyncratic error terms, conditional on the individual effect, are uncorrelated with past, present and future values of the regressors. If this fails, then all the classic panel data methods and specification tests are inconsistent. Formally, the strict exogeneity assumption means: $E(\varepsilon_{it}|Z_{is}, c_i) = 0$, for all t and s. There are reasons to suspect that the assumption might fail, as any pricing error (ε_{it}) could affect the future values of certain indicators, like reserves, debt to GDP, participation on the market, proportion of countries in arrears.

The strict exogeneity assumption is even more problematic and crucial in the structural form (equation (1) with the country effects included) than in the reduced form. As Keane and Runkle (1992) strongly point out, in this type of models, there are never any strict exogenous variables or instruments. This formal result comes from the effect of the prediction error on the future values of the variables.

When there are concerns about strict exogeneity, the general approach is to use a transformation to remove the country effects c_i, and then search for instrumental variables, assuming only sequential exogeneity (Wooldridge, 2002). According to this assumption, the idiosyncratic errors, conditional on c_i, should be uncorrelated with the contemporaneous and past values of the regressors (instruments), but not with future values.

In this respect, a first-difference (FD) estimator is attractive:

$$s_{it} - s_{i(t-1)} = \beta(R_t - R_{t-1}) + \Gamma(Z_{it} - Z_{i(t-1)}) + \varepsilon_{it} - \varepsilon_{i(t-1)}.$$

One can notice that if strict exogeneity fails, then there is a problem here as well, since $E(\varepsilon_{i(t-1)}|Z_{it}) \neq 0$. The sequential exogeneity assumption, however, implies that all the lags of Z (or their linear combination) can be used as potential instruments for $Z_{it} - Z_{i(t-1)}$ and then the estimation is consistent. In practice, however, this bias turned out to be negligible, but it decreased the precision of the estimates. For this reason, we report estimates without this extra instrumenting step.

For the structural form, this would mean estimating:

$$s_{it} - s_{i(t-1)} = \beta(R_t - R_{t-1}) + \lambda(d_{it} - d_{i(t-1)}) + \Theta(X_{it} - X_{i(t-1)}) - \lambda(\varepsilon_{2it} - \varepsilon_{2i(t-1)}) + \varepsilon_{1it} - \varepsilon_{1i(t-1)}.$$

However, the first difference specification causes further complications. The rational expectation assumption guarantees that the prediction error ε_{2it} is orthogonal to time t information, but this is not true about $\varepsilon_{2i(t-1)}$. The remedy is to use $Z_{i(t-1)}$ or $Z_{i(t-2)}$ as instruments, as those variables are not correlated with any error at time t or $t-1$.

Appendix C: Table not for publication

Table 5: Robustness of the structural-form results (not for publication)

	Specification		
	(1)	(2)	(3)
Future default	0.291	0.262	0.326
	(1.83)**	(1.7)**	(1.84)**
Benchmark yield	−0.109	−0.108	−0.109
	(−11.91)**	(−12.13)**	(−12.05)**
US BAA spread	0.233	0.23	0.234
	(7.97)**	(7.66)**	(7.93)**
Recent default	7.22	7.19	7.18
	(1.75)**	(1.72)**	(1.76)**
First stage relevance:			
Partial R^2 for future default	0.254	0.254	0.2391
Kleibergen-Paap rk Wald stat p-val	0.006	0.006	0.004
Structural form:			
Overidentification test p-value	0.742	0.69	0.744
Number of observations	151	151	151

Column 1 is our benchmark specification. Column 2 uses only quantities as weights for the spreads. Column 3 includes a 3 year grace period for computing the future default.