Market Expectations and Default Risk Premium in Credit Default Swap Prices: *A Study of Argentine Default*

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In the last several years, the credit default swap market has experienced explosive growth. A 2007 survey by Bank of International Settlements reported that the total notional amount of outstanding global credit default swap (CDS) contracts, including both single name and index, reached $45 trillion in June 2007. A conspicuous feature of the market is the continued dominance of the single-name CDS contracts with a total notional of $25 trillion. Over the same span of years, the credit market has also witnessed several major credit events, such as defaults by Argentina, Enron, and Russia.

Despite the increasing importance of the CDS market and the repeated default events in recent years, there have been few empirical investigations into the expectations of the credit market of an eventual default, using CDS data. Even less appreciated is the properties of default risk premiums implicit in CDS prices. It is tempting to ask the following questions: What are the market expectations before an actual default? How is default risk priced in credit derivatives? How does the term structure of default risk premiums implicit in the CDS prices evolve over time? What drives the variation in default risk premiums? Do the third-party rating agencies lead or lag the credit market in downgrading and upgrading the debts, and how do the rating-implied default probabilities compare to the ones expected by the market? These questions so far have largely been unanswered in literature.

The purpose of this article is to address these questions in the case of Argentine default. To this goal, we propose a valuation framework for CDS that allows dynamic interaction between interest rate, default arrival intensities, and expected default recovery, and at the same time accommodates counter-party default risk. We next develop a parsimonious three-factor parametric CDS model, which takes into account the effects of both economy-wide factors and name-specific distress on the pricing of CDS contracts. In the model, we relate the default arrival intensity of the reference debt to the underlying factors and adopt a flexible specification of market price of risks that is instrumental in our exploration of default risk premiums. We then implement the model to a unique data set of CDS on Argentine sovereign debt. The richness of the data set in both the cross-section and the time-series dimensions makes it possible for us to study the term structure of default risk premiums and how it evolves over time.

We find that, as expected, both pseudo-physical and risk-neutral default probabilities increase dramatically over the course of the sample period when a series of political and economic developments in Argentina unfolded and pointed to a possible default.[1] One-year pseudo-physical default probability rose from an average of 6.7 percent in the period before March 2001 to about 15 percent between March and July 2001. The one-year pseudo-physical default probability then jumped to an average of 42 percent after July 2001 when massive bank runs took place throughout Argentina.

Our results show that default risk premiums implicit in CDS prices are substantial and the relation between default risk premiums and default probability is not monotone. Default risk premium, which is defined as the difference between the annualized risk-neutral and pseudo-physical default probabilities over a certain horizon divided by the annualized pseudo-physical default probability, averages about 56 percent for the one-year contract during the period before March 2001. From the start of the sample period, default risk premiums rose steadily along with ascending pseudo-physical default probability and reached its peak around October 2000 when the market realized there was a serious chance of default. After October 2000, expected default probability continued to rise, however, default risk premiums declined considerably to an average of 38 percent for the period from March to July 2001, and it further dropped to an average of 24 percent after July 2001. These results suggest that, when the default probability is low, CDS investors are willing to pay a relatively high premium to hedge default, and default risk premium plays a primary role in the valuation of credit derivatives. Default risk premium rises along with moderate increase in pseudo-physical default probability. On the other hand, when the pseudo-physical default probability becomes extremely high, the actual default probability becomes relatively more important in the pricing of the CDS contracts while the impact of risk premium declines.

We also find that the term structures of the pseudo-physical and risk-neutral default probabilities may have different shapes. The annualized pseudo-physical default probability is consistently downward-sloped throughout our sample period, which confirms previous evidence on the term structure of default probabilities of speculative debt. On the other hand, the term structure of risk-neutral default probability is upward-sloped during the bulk of the sample period and it turns downward-sloped only when default risk reaches extremely high levels at the end of the sample period. Accordingly, default risk premium displays an upward-sloped term structure during the entire sample period.

Comparing with default probabilities expected by the credit market, we find that both Moody's and Standard and Poor's lagged the market in downgrading the debt, usually by several months. The implied Argentina-specific distress factor is found to be highly correlated with the JP Morgan EMBI bond spread index for Argentina, while there is virtually no correlation between the extracted Argentina-specific distress factor and the return on the Merval stock index of Argentina.

Three recent empirical studies are complementary to our work in spirit, but each of them has a focus different from ours. Driessen [2005] estimates the default risk premium using U.S. corporate bond prices instead of CDS data, and he uses average default frequencies by credit ratings as the benchmark for pseudo-physical default probabilities. He finds the average ratio of risk-neutral to pseudo-physical default probability to be 1.89 in his sample. Berndt Douglas, Duffie, Ferguson, and Schranz [2005] examine the relation between risk-neutral and pseudo-physical default probabilities on a sample of 64 firms, using CDS data and Moody's KMV EDF measure, and find dramatic variation in risk premium over time. Both studies look at corporate names, while we study CDSs on Argentine sovereign debt. Furthermore, we focus on the whole term structure of default risk premiums and analyze the evolution of the term structure of risk premiums during different stages before the eventual default. Pan and Singleton [2005] explore the property of the risk-neutral credit-events intensities by examining the square-root, lognormal, and three-halves processes. They focus on Mexico, Turkey, and Russia in their study and their sample periods are different from this study.

Other recent empirical studies in the literature include Blanco, Brennan, and Marsh [2004], Cossin, Hricko, Aunon-Nerin, and Huang [2002], Houweling and Vorst [2005], Hull, Predescu, and White [2003] and Longstaff, Mithal, and Neis [2005], among others. Cossin, Hricko, Aunon-Nerin, and Huang investigate the determinants of CDS spread, while Houweling and Vorst examine the pricing performances of several CDS models. Blanco, Brennan, and Marsh and Longstaff, Mithal, and Neis study the valuation of CDSs relative to the cash bond market, and examine which market leads in price discovery. Hull, Predescu, and Whit analyze the relationship between CDS spreads and bond yields, and explore the extent to which credit rating announcements by Moody's are anticipated by the CDS market.[2]

CDS VALUATION

In this section, we present a valuation framework for CDS that allows dynamic interaction between interest rate, default arrival intensities, and expected default recovery and, at the same time, accommodates counter-party default risk. All these features are attractive and need to be modeled. Correlation between interest rates and credit spreads are well documented in empirical literature, while the separation of expected recovery and default probability is long desired by the trading desk. In light of the credit crisis on Wall Street that started in July 2007, counterparty default risk has become an extremely important consideration for credit derivatives traders. To our knowledge, this is the first study to incorporate all these appealing features into a single framework.

A CDS can be viewed as an insurance contract in which the CDS buyer makes periodic premium payments to the seller until the maturity of the contract or default by the underlying debt, whichever occurs first. In return, the CDS buyer can ship the defaulted debt to the seller for the face value of the debt if default occurs before maturity of the CDS contract. In our framework, we allow default by the CDS seller besides default by the underlying debt. So we have a situation similar to a first-to-default credit event basket, which features valuation of contingent claims whose payoff depends not only on the timing of the first credit event, but also on the identity of the first event. We make the simplifying assumption that the CDS premium is paid continuously.

We fix a probability space $(\Omega, \mathcal{G}, Q)$, with filtration $\mathcal{G} \equiv \{\mathcal{G}_t \mid 0 < t < T\}$ satisfying $\mathcal{G}_T = \mathcal{G}$ that is complete, increasing, and right continuous, where Q is the equivalent martingale measure. We also take as given a "locally" risk-free process r. Let $\chi_1(t) = 1_{\xi_1 \le t}$ be the default indicator function of the underlying reference and $x_2(t) = 1_{\xi_2 \le t}$ be the CDS seller default indicator, where ξ_1 and ξ_2 are respectively the stopping times that characterize time of default by the underlying reference and by the CDS seller. The relevant stopping time of the first-to-default credit event basket is $\xi = \min\{\xi_1, \xi_2\}$, with corresponding credit event indicator function $\chi(t) = 1_{\xi \le t}$.

For a plain vanilla binary CDS, there are two "legs": The premium leg (i.e., the stream of CDS premiums) and the default protection leg. The present value of the premium leg is:

$$E^Q\left\{\int_t^{t+\tau} \frac{B(t)}{B(u)}\left(1-\chi(u_-)\right) p_\tau \, du \Big| \mathcal{G}_t\right\} \tag{1}$$

where $B(t) \equiv e^{\int_0^t r(u)du}$ is the money market account and p_τ is the continuous premium paid by the buyer for the CDS contract with maturity τ. The expectation is taken under the equivalent martingale measure Q.

For the default protection leg, we make the conventional assumption that, if the first credit event (before the expiration of the CDS contract) happens to be default by the underlying reference, the CDS buyer will get payoff w_1 from the CDS seller for each unit face value of the underlying debt. If the CDS seller defaults on its own debts before default by the underlying reference, the CDS contract will terminate with the CDS buyer receiving no payment (i.e., $w_2 = 0$) from the seller.[3] The payoff process of the default protection is (see Duffie [1999b]

$$\begin{aligned} dD(t) &= (1-\chi(t\_))[w_1 d\chi_1(t) + w_2 d\chi_2(t)] \\ &= (1-\chi(t\_))y_1(t)h_1(t)dt + dM_D(t) \end{aligned}$$

where $M_D(t)$ is a martingale with respect to Q, $y_1(t)$ can be viewed as the risk-neutral expected payment, conditional on all information up to but not including time t, that is $y_1(\mathrm{t}) = E^Q[w_1 \mid \mathcal{G}_{t-}]$, and $h_1(t)$ is the instantaneous intensity of default arrival of the underlying debt at time t. The present value of the payoff at default from the protection leg can then be expressed as

$$E^Q\left\{\int_t^{t+\tau} \frac{B(t)}{B(u)}\left(1-\chi(u_-)\right) y_1(u) h_1(u) du \,\Big|\, \mathcal{G}_t\right\} \tag{2}$$

By the fact that the net present value of a CDS at its initiation is zero, the fair-value CDS premium can be obtained by equating the values of the two legs,

$$p_\tau = \frac{E^Q\left\{\int_t^{t+\tau} \frac{B(t)}{B(u)}\left(1-\chi(u_-)\right) y_1(u) h_1(u) du \,\Big|\, \mathcal{G}_t\right\}}{E^Q\left\{\int_t^{t+\tau} \frac{B(t)}{B(u)}\left(1-\chi(u_-)\right) du \,\Big|\, \mathcal{G}_t\right\}} \tag{3}$$

Given the assumption that there is zero probability that both the CDS seller and the underlying reference of the CDS default at exactly the same instant of time, Equation (3) can be expressed as (see Duffie [1999b]):

$$p_\tau = \frac{E^Q\left\{\int_t^{t+\tau} y_1(u) h_1(u) e^{-\int_t^u (r(s)+h_1(s)+h_2(s))ds} du \Big| \mathcal{G}_t\right\}}{E^Q\left\{\int_t^{t+\tau} e^{-\int_t^u (r(s)+h_1(s)+h_2(s))ds} du \Big| \mathcal{G}_t\right\}} \tag{4}$$

Equation (4) provides a flexible framework that allows interaction between interest rate process, default arrival intensities, and default recovery and, at the same time, accommodates counter-party default risk. The expected recovery rate of the underlying reference, γ_1, can be specified as a function of the state variables in the model. For ease of implementation, it can also be specified as a constant.

A PARAMETRIC CDS MODEL

In this section, we present a three-factor CDS model, which allows flexible interaction between interest rate and hazard rate processes. The model is adapted from the standard reduced-form approach such as Duffie and Singleton [1999] and Jarrow and Turnbull [1995].

Following Pearson and Sun [1994] and Duffee [1999], we first specify the instantaneous default free interest rate process as the sum of a constant and two economy-wide stochastic variables, $s_1(t)$ and $s_2(t)$, that each follows a squared-root process:

$$r(t) = \alpha_r + s_1(t) + s_2(t) \tag{5}$$

$$ds_i(t) = \kappa_i(\theta_i - s_i(t))dt + \sigma_i\sqrt{s_i(t)}dW_i(t),\ i = 1, 2 \tag{6}$$

where $W_1(t)$ and $W_2(t)$ are standard Brownian motions and are independent from each other.

We also assume that there is a name-specific distress variable, $Z(t)$, associated with the underlying reference bond, which follows the following process,

$$dZ(t) = \kappa_z(\theta_z - Z(t))\,dt + \sigma_z\sqrt{Z(t)}dW_z(t) \tag{7}$$

where $W_z(t)$ is a standard Brownian motion independent from $W_1(t)$ and $W_2(t)$. This name-specific distress variable can be viewed as representing the name-specific component of default risk that is closely correlated to the financial distress of the borrower (see Bakshi, Madan, and Zhang [2006]).

The underlying state variable set in this economy is, $X(t) \equiv (s_1(t), s_2(t), Z(t))'$. Such a three-factor structure is not "maximal" in the sense of Dai and Singleton [2000], since the mean-reversion matrix of the state variable vector is not a full matrix. However, this model gives us a right balance between tractability and adequacy, and it has been used extensively in the credit risk literature (see Duffee [1999]). As we will show later on, given a flexible market price of risks specification, this model satisfactorily explains the U.S. Treasury term structure and captures the name-specific default risk of the underlying entity.

We consider a flexible specification for the market price of risks in the underlying state variables,

$$\gamma(X_t) = \gamma_0(t) + \gamma_1(t)X_t \tag{8}$$

where $\gamma_0(t)$ is a 3×1 vector and $\gamma_1(t)$ is a 3×3 diagonal matrix. This specification of market price of risks allows risk adjustment on both the level and the strength in the mean-reversion in state variables. The state variables in the economy then follow the following dynamics under the equivalent martingale measure,

$$\begin{aligned} ds_i(t) &= \left[\kappa_i\theta_i + \delta_i - (\kappa_i + \lambda_i)s_i(t)\right]dt \\ &\quad + \sigma_i\sqrt{s_i(t)}\,d\tilde{W}_i(t),\ i = 1, 2 \\ dZ(t) &= \left[\kappa_z\theta_z + \delta_z - (\kappa_z + \lambda_z)Z(t)\right]dt \\ &\quad + \sigma_z\sqrt{Z(t)}\,d\tilde{W}_z(t) \end{aligned}$$

where $\tilde{W}_1$, $\tilde{W}_2$, and $\tilde{W}_3$ are independent standard Brownian motions under the equivalent martingale measure Q.

Following Duffee [1999], we make the convenient assumption that the hazard rate of the underlying reference bond, $h_1(t)$, is linear in the three state variables in the economy:

$$h_1(t) = \Lambda_0 + \Lambda_{s_1}(s_1(t) - \bar{s}) + \Lambda_{s_2}(s_2(t) - \bar{s}_2) + Z(t) \tag{9}$$

where parameters Λ_{s1} and Λ_{s2} reflect correlation between the hazard rate and the interest rate. $\bar{s}_1$and $\bar{s}_2$ are the time-series means of $s_1(t)$ and $s_2(t)$.

Since most CDS sellers are big financial institutions whose financial conditions are not directly linked to a particular debt, it is reasonable to assume that the hazard rate of the CDS seller, $h_2(t)$, does not depend on the name-specific distress variable Z, rather it is a linear function of the two economy-wide factors,

$$h_2(t) = \varphi_0 + \varphi_{s_1}(s_1(t) - \bar{s}_1) + \varphi_{s_2}(s_2(t) - \bar{s}_2) \tag{10}$$

Observe from Equations (4) and (10) that our CDS model will collapse into the case of no counter-party default if the default intensity process of the CDS seller, $h_2(t)$, is zero. To keep the model parsimonious, we assume that the conditional expected loss at default, γ_1, does not depend on any of the three state variables.

The dynamics of the state variables, as specified in Equations (5) to (7), and the hazard rate specifications (plus the default recovery) completely determine the valuation process of the financial securities in our CDS model. Following the ideas of Bakshi and Madan [1999] and Duffie, Pan, and Singleton [1999], we define the following characteristic function.

$$\Phi(t,\tau;\phi) \equiv E_t^Q\left[e^{-\int_t^{t+\tau}\left[r(u)+h_1(u)+h_2(u)\right]du+i\phi h_1(t+\tau)}\right] \quad (11)$$

subject to the boundary condition, $\Phi(t+\tau, 0; \phi) = e^{i\phi h_1(t+\tau)}$. The following proposition gives the analytical solution of this characteristic function $\Phi(t,\tau,\phi)$ and the CDS premium expressed in terms of $\Phi(t,\tau,\phi)$ (see the Appendix for proof).

Proposition 1. *Let the interest rate process follow (5)–(6), name-specific distress factor follow (7), and default arrival intensities for the underlying reference and CDS seller be of (9) and (10). Given the characteristic function, $\Phi(t,\tau;\phi)$, defined as in (11), we have:*

1. *The characteristic function $\Phi(t,\tau;\phi)$ can be analytically solved as:*

$$\Phi(t,\tau;\phi) = e^{A(t,\tau;\phi)-B(t,\tau;\phi)s_1(t)-C(t,\tau;\phi)s_2(t)-D(t,\tau;\phi)Z(t)} \quad (12)$$

with

$$\begin{aligned} A(t,\tau;\phi) &= A_1(t,\tau;\phi)+A_2(t,\tau;\phi)+A_3(t,\tau;\phi) \\ &\quad + i\phi(\Lambda_0-\Lambda_{s_1}\bar{S}_1-\Lambda_{s_2}\bar{S}_2)-(\alpha_r+\Lambda_0 \\ &\quad +\varphi_0-\Lambda_{s_1}\bar{S}_1-\Lambda_{s_2}\bar{S}_2-\varphi_{s_1}\bar{S}_1-\varphi_{s_2}\bar{S}_2)\tau \end{aligned}$$

$$B(t,\tau;\phi) = \frac{-i\phi\Lambda_{s_1}\left[\gamma_1\coth\left(\frac{\gamma_1\tau}{2}\right)-(\kappa_1+\lambda_1)\right]+2\left(1+\Lambda_{s_1}+\varphi_{s_1}\right)}{\gamma_1\coth\left(\frac{\gamma_1\tau}{2}\right)+\left[(\kappa_1+\lambda_1)-i\phi\Lambda_{s_1}\sigma_1^2\right]}$$

$$C(t,\tau;\phi) = \frac{-i\phi\Lambda_{s_2}\left[\gamma_2\coth\left(\frac{\gamma_2\tau}{2}\right)-(\kappa_2+\lambda_2)\right]+2(1+\Lambda_{s_2}+\varphi_{s_2})}{\gamma_2\coth\left(\frac{\gamma_2\tau}{2}\right)+\left[(\kappa_2+\lambda_2)-i\phi\Lambda_{s_2}\sigma_2^2\right]}$$

$$D(t,\tau;\phi) = \frac{-i\phi\left[\gamma_3\coth\left(\frac{\gamma_3\tau}{2}\right)-(\kappa_z+\lambda_z)\right]+2}{\gamma_3\coth\left(\frac{\gamma_3\tau}{2}\right)+\left[(\kappa_z+\lambda_z)-i\phi\sigma_z^2\right]}$$

where $A_1(t, \tau; \phi) - A_3(t, \tau; \phi)$, and $\gamma_1 - \gamma_3$ are provided in the Appendix.

2. *Given the characteristic function in* (12), *the CDS premium in* (4) *can be expressed as*

$$p_\tau = \frac{\gamma_1\int_t^{t+\tau}\frac{1}{i}\frac{\partial\Phi(t,u;\phi)}{\partial\Phi}\Big|_{\phi=0}du}{\int_t^{t+\tau}\Phi(t,u;\phi=0)du} \quad (13)$$

where $\Phi(t, u; \phi=0)$ and $\partial\Phi(t,u;\phi)/\partial\phi|_{\phi=0}$ are respectively, $\Phi(t, u; \phi)$ and the derivative of $\Phi(t, u; \phi)$ with respect to ϕ evaluated at $\phi = 0$, whose expressions are given in the Appendix.

The above proposition shows that the characteristic function defined in Equation (11) synthesizes the problem of CDS valuation and all relevant security prices can be obtained by manipulating this characteristic function. In the remainder of this article, we implement the parametric model to a unique sample of CDSs on Argentine sovereign debt from January 1999 to December 2001. Our empirical investigation focuses on the expectations of the credit market about the default prospect of Argentine debts and the term structure of default risk premiums implicit in CDS prices.

DATA AND ESTIMATION STRATEGY

The raw data of CDSs on Argentine sovereign debt used in our study include daily closing mid-market quotes from JP Morgan's trading desk on 10 contracts with maturities ranging from 1 to 10 years. The CDS contracts and their underlying reference debts are all quoted in U.S. dollars, which eliminates the complication of exchange rate in the valuation process. The CDS premium is quoted as an annual percentage of the notional amount. The data sample covers the period from January 14, 1999, to December 05, 2001, with 749 daily observations and a total of 7,490 CDS quotes. The advantage of the data set is that there is a "True" or "False" flag for each observation, indicating whether the quotes on a particular day were true quotes from JP Morgan or some stalled quotes. So we delete all observations with a "False" flag. This screening leaves us with 698 daily observations and a total of 6,980 CDS quotes. To reduce noise from the daily observations, we construct the corresponding weekly data series from the daily series of CDS quotes by picking observations on each Wednesday. The resulting sample includes 151 weekly observations from January 20, 1999, to December 05, 2001, with a total of 1,510 price quotes. This weekly data sample provides the basis for our empirical analysis that follows.

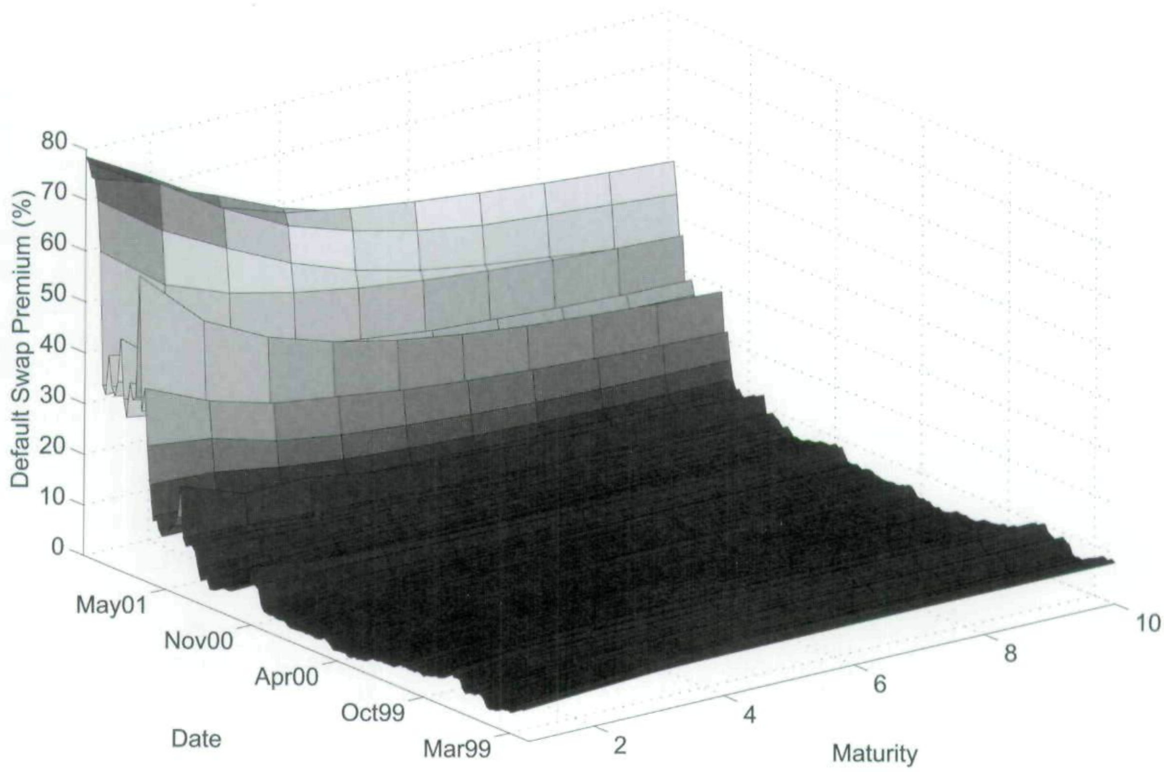
Exhibit 1 shows the evolution of the term structure of the premium of CDSs on Argentine sovereign debts over the sample period. For the majority of the sample period before March 2001, the CDS premiums were in the range of 3 to 9 percent. An exception was a brief blip in November 2000, when the premium on short contracts jumped over 10 percent. However, since mid-March of 2001, the CDS premium on short contracts spiked to the magnitude of over 10 percent and further jumped to 30 to 40 percent in mid-July 2001, and eventually reached the 60 to 70 percent level. On the other hand, the slope of the term structure of default swap premium was upward-sloping before March 2001, except the brief reversal in November 2000, but it turned into downward-sloping after March 2001.

A series of economic and political events took place before the eventual default by the Argentine government.[4] Two of them stand out in their significance. The first is the rejection of Lopez Murphy's austerity plan in March of 2001, which signaled the end of any realistic chance that the Argentine government had the political resolve and power to achieve the fiscal discipline needed to avoid a sovereign default. Another event was the large scale withdrawals of deposit from Argentine banks in late June and early July of 2001, which led to eventual total collapse of the financial system and market confidence. For this reason, we split the whole sample period into three sub-periods: 1) the normal period—from the start of the sample period to 03/14/2001; 2) the transition period—from 03/21/2001to 06/27/2001; and 3) the crisis period—from 07/05/2001 to the end of the sample period.

Exhibit 2 shows summary statistics of the CDS premium on Argentine sovereign debt over the three sub-periods. A few observations are in order. First, the average premiums of CDS rates are vastly different over the three sub-periods. Second, the term structure of the CDS premium is upward-sloping in the normal period, while

EXHIBIT 1

Credit Default Swap Prices on Argentine Sovereign Debt (01/20/1999–12/05/2001)



it is downward-sloping in the transition period and even more so in the crisis period. Third, as shown by the standard deviation, there are more variations at the short end of the term structure of CDS premium.

Estimation Strategy

In the empirical estimation, we implement a two-step procedure to estimate the parameters of the term structure of interest rate and that of the CDS. In the first step, we estimate the parameters of the term structure of U.S. interest rate, using the extended Kalman filter; and in the second step, we estimate the remaining parameters related to CDS process using a quasi-maximum likelihood (QML) method, taking parameters of term structure of interest rate and the two macroeconomic state variables estimated in the first step as inputs.[5]

Estimation of Term Structure of U.S. Interest Rates. In the estimation of the U.S. interest rate process, we use the plain-vanilla fixed for float U.S.-dollar LIBOR-quality interest rate swap yields as the basis for the U.S. term structure of interest rate (see Dai and Singleton [2000] and Duffie, Pedersen, and Singleton [2003]).[6] Under the assumption of interest rate swaps being default-free, the fair-value swap rate with maturity τ_m at its initiation is (see Duffie and Singleton [1997]):

$$c_t^m = \frac{1 - B_0(t, \tau_m)}{\sum_{j=1}^{2\tau_m} B_0\left(t, \frac{j}{2}\right)} \tag{14}$$

where $B_0(t, \frac{j}{2})$ is the risk-free zero coupon bond price with time to maturity $\frac{j}{2}$. In our parametric CDS model, the risk-free zero-coupon bond price $B_0(t, \frac{j}{2})$ is the two-factor extended CIR bond price with time to maturity $\frac{j}{2}$.

We then estimate the parameters of the U.S. term structure of interest rates, using the extended Kalman filtering (see Geyer and Pichler [1996] and Duffee [1999]). The transition equation in an approximate Kalman filter in combination with QML is,

$$s(t) = g_t + F_t s(t-1) + u(t) \tag{15}$$

EXHIBIT 2
Summary Statistics of Default Swap Premiums on Argentine Sovereign Debt

This exhibit summarizes the weekly credit default swap premia on Argentine sovereign debt from 01/20/1999 to 12/05/2001. The sample period is split into three sub-periods: 1) the normal period (01/20/99–03/14/01); 2) the transition period (03/21/01–06/27/01); and 3) the crisis period (07/05/01–12/05/01). For each of these three sub-periods, the mean, standard deviation, skewness, and excess kurtosis of default swap premium with maturities from 1 to 10 years are reported. Nobs - Numbers of weekly observations. Data source: JP Morgan Chase.

The Normal Period (01/20/99–03/14/01)

	1 yr	2 yr	3 yr	4 yr	5 yr	6 yr	7 yr	8 yr	9 yr	10 yr	Nobs
Mean	4.350	5.214	5.653	5.997	6.246	6.418	6.550	6.648	6.725	6.786	111
SDev	1.711	1.567	1.458	1.351	1.310	1.310	1.319	1.334	1.347	1.359	
Skew	1.139	0.598	0.468	0.440	0.417	0.370	0.344	0.333	0.339	0.354	
Kurt	2.133	-0.188	-0.604	-0.763	-0.833	-0.854	-0.795	-0.697	-0.604	-0.515	

The Transition Period (03/21/01–06/27/01)

	1 yr	2 yr	3 yr	4 yr	5 yr	6 yr	7 yr	8 yr	9 yr	10 yr	Nobs
Mean	12.530	11.716	11.737	11.216	11.053	10.752	10.603	10.393	10.285	10.280	15
SDev	2.987	1.798	2.023	1.512	1.538	1.495	1.301	1.459	1.444	1.179	
Skew	2.102	2.142	1.555	1.450	0.827	0.598	0.340	0.289	0.232	0.031	
Kurt	4.149	4.527	2.364	2.162	0.196	-0.264	-0.748	-0.723	-0.771	-1.105	

The Crisis Period (07/05/01–12/05/01)

	1 yr	2 yr	3 yr	4 yr	5 yr	6 yr	7 yr	8 yr	9 yr	10 yr	Nobs
Mean	45.688	41.216	36.835	34.026	32.116	30.743	29.737	28.972	28.362	27.896	23
SDev	18.308	16.124	14.953	13.557	12.673	12.108	11.645	11.303	10.973	10.682	
Skew	0.755	0.915	0.906	0.888	0.898	0.916	0.916	0.921	0.932	0.929	
Kurt	-0.900	-0.584	-0.598	-0.566	-0.491	-0.417	-0.383	-0.363	-0.328	-0.310	

where $s(t) = (s_1(t), s_2(t))'$, $E_{t-1}[u(t)] = 0$, and $E_{t-1}[u(t)u(t)'] = R_t$ is a 2×2 diagonal matrix with j-th element given by $\sigma_i^2 k_i^{-1}(1-e^{-k_j\Delta t})[\frac{\theta_j}{2}(1-e^{-k_j\Delta t})+e^{-k_j\Delta t}s_j(t-1)]$. g_t is a 2×1 vector with j-th element $\theta_j(1 - e^{-\kappa j\Delta t})$, and F_t is a 2×2 diagonal matrix with (j, j) element, $e^{-\kappa j\Delta t}$.

At each week, let the observed interest rate swap rates $Y(t) = (y_1(t),\ldots,(y_n(t))'$, where n is the number of swap rates to be used in the estimation,

$$Y(t) = Y(s(t)) + \varepsilon(t),\ t = 1,\ \ldots,\ T \tag{16}$$

where $Y(s(t))$ is an $n \times 1$ vector with i-th element $y_i(s(t),\tau_i) = 1 - B_0(t,\tau_i)/\Sigma_{k=1}^{2\tau_i} B_0(t,\frac{k}{2})$, and $B_0(t,\frac{k}{2})$ is the risk-free zero coupon bond price at time t with maturity $\frac{k}{2}$. $\varepsilon(t) \sim N(0, H_t)$, where H_t is an $n \times n$ diagonal matrix with element $H_{i,i} = \Sigma_i$ Linearizing the measurement equation at $s(t) = \hat{s}(t\,|\,t-1)$, we have the following approximate measurement equation for the i-th element,

$$\begin{aligned} Y_i(t) &= y_i\left(\hat{s}(t|t-1),\tau_i\right) + \frac{\partial y_i(s(t))}{\partial S'(t)}\Big|_{s=\hat{s}(t|t-1)} \\ &\quad \times\left[s(t) - \hat{s}(t|t-1)\right] + \varepsilon_i(t) \\ &= a_i(t) + b_i(t)s(t) + \varepsilon_i(t) \end{aligned} \tag{17}$$

where $a_i(t) = [y_i(\hat{s}(t\,|\,t-1),\tau_i) - \frac{\partial_{y_i}(s(t))}{\partial s'(t)}\big|_{s=\hat{s}(t|t-1)}\hat{s}(t|t-1)]$ and $b_i(t) = \frac{\partial_{y_i}(s(t))}{\partial s'(t)}\big|_{s=\hat{s}(t|t-1)}$.

Standard Kalman iteration gives the log-likelihood function

$$L = -\frac{nT}{2}\log(2\pi) - \frac{1}{2}\sum_{t=1}^{T}\log\left|M_t\right| - \frac{1}{2}\sum_{t=1}^{T}(\upsilon_t' M_t^{-1}\upsilon_t) \tag{18}$$

where v_t are pricing errors that are independently and normally distributed with mean zero and a covariance matrix, $M_t = b_t P_{t|t-1|} b_t' + H_t$, with $P_{t|t-1} = F_t P_{t-1} F_t' + R_t$. This log-likelihood function can be maximized to generate the parameter estimate of the state-space system.[7]

Estimation of CDS Parameters. Taking parameters of U.S. interest rates estimated in the first step as inputs, along with the estimated time-series of the two economy-wide state variables, $(\hat{s}_1(t), \hat{s}_2(t))$, we estimate the remaining parameters related to the default swaps using a standard QML method in the second step.

Specifically, we assume that the premium on the 5-year CDS, which are usually the most liquid contracts, is measured without error. Given the parameter set and the two estimated economy-wide state variables $\hat{s}_1(t)$ and $\hat{s}_2(t)$, the implied process for Argentina-specific distress factor, $Z(t)$, can be inverted from the 5-year CDS formula given in Equation (13). The density function of the 5-year default swap premium c_t^5 conditional on c_{t-1}^5 can be expressed as $\frac{1}{|J_t^C|} f_z(Z_t\,|\,Z_{t-1})$, where J_t^C is the corresponding Jacobian of the transformation at time-t that is non-linear and time-dependent. As we know, the conditional densities of the state variables $f_z(Z|\,Z_{t-1})$ are non-central chi-square, as shown in Cox, Ingersoll, and Ross [1985].

For the CDS contracts with maturities of 1, 3, and 10 years, the CDS premiums are assumed to be measured with errors. We assume that the nonzero measurement errors $\{\varepsilon_t\}$of 1-, 3-, and 10-year CDS contracts are serially uncorrelated, but jointly normally distributed with zero mean and variance-covariance matrix Ω_ε (see Duffee [2002]).

Under these assumptions, the log-likelihood function for a sample of observations on the 5-year CDS premium, and the three imperfectly-observed (1-, 3-, and 10-year) CDS yields for $t = 2, \ldots, T$, in the conditional maximum likelihood estimation is,[8]

$$\begin{aligned} L &= \sum_{t=2}^{T}\log f_Z\left(\hat{Z}_t\middle|\hat{Z}_{t-1}\right) - \sum_{t=2}^{T}\log\left|J_t^C\right| \\ &\quad - \frac{3(T-1)}{2}\log(2\pi) - \frac{T-1}{2}\log\left|\Omega_\varepsilon\right| \\ &\quad - \frac{1}{2}\sum_{t=2}^{T}\varepsilon_t'\Omega_\varepsilon^{-1}\varepsilon_t \end{aligned} \tag{19}$$

where expressions for $f_z(Z_t\,|\,Z_{t-1})$ and J_t^C are given in the Appendix.

For the purpose of implementing the QML method, we substitute the exact transition density $f_z(Z_t\,|\,Z_{t-1})$ with a normal density: $Z(t)\,|\,Z(t-1) \sim N(\mu_t, Q_t)$, where μ_t and Q_t are the first two moments of $Z(t)$ given $Z(t-1)$. To ensure the variance-covariance matrix Ω_ε of the pricing errors of the 1-, 3-, and 10-year CDS contracts be invertible, we assume that Ω_ε, which is time-invariant, satisfies the Cholesky decomposition, $\Omega_\varepsilon = CC'$, where C is a 3×3 matrix with non-zero elements C_{11}, C_{22}, C_{33}, C_{21}, and C_{32}. The advantage of our empirical procedure is that both cross-sectional and time series information is used in estimation, which makes it possible to separately identify the time-series parameters of the state variables and that of the market price of risks.

EMPIRICAL RESULTS

In this section, we first discuss the parameter estimation that includes estimation of U.S. interest rate process and that of the CDS. We then look at the in-sample fit of the model. We further analyze the market expectations of the default prospect of Argentine sovereign debt and default risk premiums implied in default swap rates. Based on the calculated implied default probabilities, we examine whether Moody's and S&P led or lagged the market in downgrading Argentine debt during the sample period.

Estimation of U.S. Term Structure of Interest Rate

In the extended Kalman filtering estimation of default-free interest rate process, we select a broad cross-section of U.S. interest rate swap rates with maturities of 2, 3, 5, 7, and 10 years. For the purpose of convergence of the estimation, we impose the restriction, $\delta_2 = 0$, and we also set $\alpha_r = -0.99$ (see Duffee [1999, 2002] for similar treatment).

Panel A of Exhibit 3 reports estimation results of the two-factor interest rate model as well as the in-sample fit. The parameter estimates are intuitive and are in line with expectations. Consistent with previous studies in literature, our results show that there is strong mean reversion for the first economy-wide factor, s_1, but a rather weak mean-reversion in the second economy-wide factor, s_2. Specifically, the estimates of $k_1 = 0.8126$ and $k_2 = 0.0909$ imply a half-life of 0.853 years and 7.625 years, respectively, for s_1 and s_2 processes.

The estimates of $\lambda_1 = -0.0104$ and $\lambda_2 = -0.0724$ indicate negative risk adjustment on the mean-reversion parameter for both economy-wide factors. However, both $k_1 + \lambda_1$ and $k_2 + \lambda_2$ are positive, which shows that both s_1 and s_2 still show mean-reversion behavior under the risk-neutral measure, but at a slower pace and more persistently than under the physical measure. Estimate of the constant component on the market price of risks is negative, $\delta_1 = -0.0005$. For most parameters, the parameter estimates are much higher than their standard errors, which means they are statistically significant.

We also examined the in-sample fit of the two-factor interest rate model, and the results show that the model explains the U.S. term structure of interest rates very well. The near zero median pricing errors across all maturities suggest that there is virtually no systematic bias in the model when it is fitted to the interest swap rates. The mean absolute pricing errors are about 5 basis points (bps), and the squared-roots of the mean squared pricing errors are a little over 7 bps across all maturities. Our estimation results compare favorably with previous studies (e.g., Duffee [1999] and Babbs and Nowman [1999]). The estimated standard deviation of the measurement errors, for the 2-, 3-, 5, 7-, and 10-years are, respectively, 1.4, 0.7, 0.7, 0.5, and 1.3 bps. These results indicate the model fits the U.S. term structure of interest rates very well.

Consistent with previous studies, correlation exercises show that the first extracted factor from the extended Kalman filtering, $\hat{s}_1$, is highly negatively correlated with the slope of the U.S. Treasury yield, defined as the difference between the 10-year and the 0.5-year Treasury yields, with a correlation coefficient of –0.95. On the other hand, the second extracted factor, $\hat{s}_2$, is highly correlated with the 10-year Treasury yield with a correlation coefficient of 0.90. So the two-factor interest rate model captures the slope and the level of U.S. Treasury term structure satisfactorily.

Estimation of Default Parameters and In-Sample Fit

Using the parameter estimates for the U.S. term structure of interest rates and the two extracted economy-wide factors as inputs, we estimate the parameters of the CDS process in the second step. In our numerous estimating efforts, the model seems to have difficulty pinning down the counter-party default probability of the CDS seller. So we set both φ_1 and φ_2 to be zero and φ_0 to be a constant at 1.20% that is comparable to the risk-neutral default probability of an AA to A-rated corporate borrower.

Panel B of Exhibit 3 provides the parameter estimates of the CDS model in the QML estimation and their asymptotic standard errors using 1-, 3-, 5-, and 10-year CDS data from 01/20/1999 to 03/14/2001. The estimate of the mean-reversion factor, $\kappa_z = 0.52$, shows that the mean-reversion in Argentina-specific distress factor is relatively weak with a half-life of 1.33. Together with the estimate of a large negative λ_z, the mean-reversion parameter under risk-neutral measure, $\kappa_z + \lambda_z = -0.032$, which shows that Argentina-specific distress processes

EXHIBIT 3

Parameter Estimates of U.S. Interest Rate Process and Default Process

Panel A reports parameter estimates of the U.S. interest rate process using the extended Kalman filter. Under the risk-neutral measure, the dynamics of the two-factor interest rate model are:

$$r(t) = \alpha_\tau + s_1(t) + s_2(t)$$
$$ds_i(t) = [\kappa_i\theta_i + \delta_i - (\kappa_i + \lambda_i)s_i(t)]dt + \sigma_i\sqrt{s_i(t)}d\tilde{\omega}_r(t),\ i = 1, 2$$

and the dynamics under the physical measure are specified in Equation (6). The estimation uses weekly time-series of U.S. interest rate swap rates with maturities of 2, 3, 5, 7, and 10 years. To get convergence in estimation, α_r and δ_2 are set, respectively, to –0.99 and zero. The asymptotic standard errors are reported in parentheses. The last column reports the maximized log-likelihood function $\mathcal{L}$. Panel B reports the parameter estimates of the CDS process in the QML estimation. This estimation procedure uses four weekly CDS rates with maturity of 1, 3, 5, and 10 years from 01/20/1999 to 03/14/2001. The 5-year contract is assumed to be priced without error. The remaining three contacts are assumed to be priced with errors, where the errors are assumed to be normally distributed and serially uncorrelated but cross-sectionally correlated. The log-likelihood function to be maximized is:

$$L = \sum_{t=2}^{T} \log f_Z\left(\hat{Z}_t \middle| \hat{Z}_{t-1}\right) - \sum_{t=2}^{T} \log\left|J_t^C\right| - \frac{3(T-1)}{2}\log(2\pi) - \frac{T-1}{2}\log\left|\Omega_\varepsilon\right| - \frac{1}{2}\sum_{t=2}^{T} \varepsilon_t'\Omega_\varepsilon^{-1}\varepsilon_t \tag{20}$$

where J_t^C represents the Jacobian of the 5-year CDS rate. Asymptotic standard errors based on the robust QML estimates proposed by White [1982] are reported in the parentheses. The estimated log-likelihood value is 1735.12. Data source: Bloomberg, Federal Reserve Board and JP Morgan Chase.

Panel A: U.S. Risk free Interest Rate Parameter Estimates

Parameter	κ_i	σ_i	θ_i	λ_i	δ_i	$\mathcal{L}$
s_1	0.8126 (0.0003)	0.0129 (0.0001)	1.0033 (0.0002)	-0.0104 (0.0001)	-0.0005 (0.0003)	5226.44
s_2	0.0909 (0.0617)	0.0534 (0.0031)	0.0210 (0.0137)	-0.0724 (0.0594)	0	

Panel B: CDS Parameter Estimates

κ_z	σ_z	θ_z	λ_z	δ_z	w_1	Λ_0	Λ_{s_1}	Λ_{s_2}
0.5209 (0.0009)	0.2436 (0.0002)	0.1320 (0.0003)	-0.5529 (0.0004)	-0.0236 (0.0001)	0.4035 (0.0004)	-0.0976 (0.0007)	0.9299 (0.0008)	-0.2446 (0.0005)
C_{11}	C_{22}	C_{33}	C_{21}	C_{32}				
0.0111 (0.0005)	0.0019 (0.0001)	0.0055 (0.0003)	0.0032 (0.0001)	0.0015 (0.0004)				

display no mean-reversion behavior at all under the risk-neutral measure, which is consistent with previous evidence on speculative grade bonds.

The estimates of the two sensitivity parameters in the hazard rate specification of the Argentine sovereign debt show that the hazard rate, $h_1(t)$, is positively related to the first term structure factor, while it is negatively related to the second factor. Since the first and the second extracted U.S. interest rate factors are shown to be, respectively, closely correlated with the negative slope and the 10-year spot rate of U.S. Treasury yield, this implies that the hazard rate of the Argentine sovereign debt is negatively related to the slope of U.S. term structure of interest rate and it is positively related to the level at the long end. This result is consistent with existing evidence that the credit spread is higher during economic downturns when the slope of the term structure is usually flat (e.g., Wu and Zhang [2007]).

The estimate of the expected rate of payoff at default for the CDS holder is 0.404, which implies a recovery rate for the underlying reference at 0.596, which is higher than historical observations but roughly in line with results reported by Pan and Singleton [2005]. The estimate of the Cholesky decomposition of the variance–covariance matrix implies that the standard deviations for the pricing errors of the 1-, 3-, and 10-years default swaps are respectively, 0.0111, 0.0037, and 0.0057, with the correlation coefficient between pricing errors of 1- and 3-year default swaps at 0.854 and of 3- and 10-years default swaps at 0.134.

We also examine the in-sample fit of our CDS model. Based on parameter estimates in Exhibit 3, we calculate the model-determined CDS premium for contracts with maturities from 1 to 10 years (except the benchmark 5-year contract that is measured without error). The pricing errors are then computed as the market prices minus the model-determined prices. The percentage pricing error is calculated as the pricing error of a contract divided by the market premium of that contract. The median pricing errors for most maturities are in the magnitude of a few basis points. The median percentage pricing errors are less than 1.6% for most contracts. In terms of absolute pricing errors, the mean absolute pricing errors and the mean absolute percentage pricing errors show that the model fits the CDS data very well for most of the maturities except the very short contracts. For most contracts, the mean absolute pricing errors are in the range of 10 to 20 bps.

On the other hand, the absolute pricing errors for the 1-year contract are relatively large, which may be due to the fact that there are larger variation in the CDS premiums at the short end of the maturity, as manifested by the high standard deviation in the CDS premiums of the 1-year contract (shown in Exhibit 2).

Market Expectations and Default Risk Premiums

A set of questions that are of particular interest to us are the expectations of the credit market on the prospect of Argentine default and the default risk premiums implied in CDS prices over the course of the sample period. In particular, what were the pseudo-physical and risk-neutral default probabilities expected by the CDS investors during the sample period? What were the differences between the risk-neutral and the physical default probabilities over the different stages of the sample period before default? How did the term structure of the default risk premiums evolve over the course of the sample period? And finally, what drives the variation in default risk premiums over time? The richness of our data set and the flexible structure of our model make it possible for us to explore answers to these important questions.

Using the parameter estimates reported in Exhibit 3, we are able to back out both the physical and risk-neutral default probabilities over different time-horizons, as well as the expected recovery rate, from the observed CDS prices. Specifically, the τ-year conditional cumulative risk-neutral default probability is calculated according to a formula, $DP = 1 - E^{Q}[e^{-\int_0^{\tau} h(u)du}]$, where $h(u)$ is the time-s risk-neutral instantaneous hazard rate, and the expectation is taken under the risk neutral measure Q (see the Appendix for details). The τ-year pseudo-physical default probability can be calculated under the physical measure. We should emphasize that the pseudo-physical default probabilities calculated under the physical measure are different from the historical default probabilities because the hazard rate in our formula is still the risk-neutral hazard rate instead of the physical hazard rate. As explained by Jarrow, Lando, and Yu [2005], physical hazard rate cannot be extracted from credit derivatives prices alone. Nonetheless, the difference between the risk-neutral and pseudophysical default probabilities shows that there is risk premium associated with time-variation in the risk-neutral hazard rate $h(t)$ that is priced in the credit market. With an abuse of terminology, we define the default risk premium as the difference between the annualized risk-neutral and pseudophysical default probabilities scaled by the annualized pseudo-physical default probability.

The first two rows (DP1 and DP2) in each panel of Exhibit 4 represent, respectively, the mean annualized pseudo-physical and risk-neutral default probabilities over the time-horizon of 1 to 10 years during the normal, transition, and crisis periods of our sample. The third row in each panel reports the mean difference between the annualized risk-neutral and pseudo-physical default probabilities for maturities of 1 to 10 years. The fourth row reports the default risk premium that measures the relative size of the difference between the two annualized default probabilities.

The following observations can be made from Exhibit 4. First, both pseudo-physical and risk-neutral default probabilities increase dramatically from the normal period to the transition period and to the crisis period. For instance, the 1-year pseudo-physical default probability rose from an average of 6.72 percent in the normal period to 15 percent in the transition period and then to 42 percent in the crisis period.

Second, annualized pseudo-physical default probability exhibits downward-sloping shapes during all three sub-periods, and it is the steepest in the crisis period. On the other hand, the annualized risk-neutral default

EXHIBIT 4

The Term Structure of Default Probabilities and Default Risk Premium

This exhibit summarizes the annualized pseudo-physical and risk-neutral default probabilities (in %) as well as the difference between the two default probabilities and default risk premium, which is defined as the difference between the annualized risk-neutral and pseudo-physical default probabilities divided by the annualized pseudo-physical default probability. The first row reports the mean pseudo-physical annualized default probabalities for maturities from 1 to 10 years. The second row presents the mean of the risk-neutral annualized default probabilities. The third row summarizes the mean difference between the two default probabilities and the fourth row shows the default risk premiums from 1 to 10 years. All statistics are reported over the three sub-periods: 1) the normal period (01/20/99–03/14/01); 2) the transition period (03/21/01–06/27/01); and 3) the crisis period (07/05/01–12/05/01). The τ-year cumulative default probability under a particular measure * is calculated as $1 - E_t^*[e^{-\int_0^\tau h_1(s)du}]$, where $h_1(s)$ is the instantaneous hazard rate of the underlying debt at time s. Data source: Federal Reserve Board and JP Morgan Chase.

The Normal Period (01/20/99–03/14/01)

	1 yr	2 yr	3 yr	4 yr	5 yr	6 yr	7 yr	8 yr	9 yr	10 yr
DP1	6.725	5.806	5.117	4.603	4.219	3.931	3.703	3.532	3.400	3.294
DP2	10.123	11.962	13.400	14.511	15.323	15.911	16.336	16.638	16.861	17.012
DP2-DP1	3.398	6.156	8.283	9.908	11.104	11.980	12.634	13.106	13.461	13.718
DRPrem	55.7	114.8	172.7	227.2	275.7	316.9	353.1	382.5	406.5	426.3

The Transition Period (03/21/01–06/27/01)

	1 yr	2 yr	3 yr	4 yr	5 yr	6 yr	7 yr	8 yr	9 yr	10 yr
DP1	15.053	12.573	10.667	9.220	8.127	7.300	6.660	6.167	5.747	5.420
DP2	20.727	22.500	23.640	24.273	24.520	24.513	24.340	24.073	23.760	23.420
DP2-DP1	5.673	9.927	12.973	15.053	16.393	17.213	17.680	17.907	18.013	18.000
DRPrem	38.1	79.7	122.7	164.7	203.3	237.5	267.3	292.2	315.5	334.0

The Crisis Period (07/05/01–12/05/01)

	1 yr	2 yr	3 yr	4 yr	5 yr	6 yr	7 yr	8 yr	9 yr	10 yr
DP1	41.704	35.191	29.878	25.648	22.335	19.722	17.657	15.974	14.604	13.491
DP2	51.130	52.030	52.139	51.591	50.596	49.278	47.800	46.243	44.678	43.157
DP2-DP1	9.426	16.839	22.261	25.943	28.261	29.557	30.143	30.270	30.074	29.665
DRPrem	23.8	50.5	78.7	106.7	133.3	157.5	179.1	198.5	215.4	229.7

probability term structure shows different shapes during the three sub-periods: It is upward-sloped during the normal period and pretty much flat during the transition period, and the curve becomes downward-sloped during the crisis period when default probabilities reach very high levels. The downward-sloping shape of the pseudo-physical default probability term structure is consistent with well-documented evidence that, for a high speculative debt, the term structure of annualized conditional default probability is downward-sloped and more so as the debt becomes more speculative (see Fons [1994] and Sarig and Warga [1989]). However, our results indicate that, compared to the case of pseudo-physical default probability, the term structure of risk-neutral default probability is upward-sloped even for speculative debts and it only turns downward-sloping when default risk becomes extremely high. This result suggests the behavior of pseudo-physical and risk-neutral default probabilities are quite different.

Third, the default risk premium implicit in CDS prices is substantial, and as default probabilities rises, the default risk premium declines. For instance, average default risk premium in the 5-year CDS price was 275 percent of the size of the pseudo-physical default probability during the same period. It declined to 203 percent in the transition period and further dropped to 133 percent during the crisis period. This pattern holds for other time horizons as well. Fourth, measured by both absolute and relative terms, the term structure of default risk premiums displays an upward-sloping shape during all three sub-periods. For instance, measured as a percentage of the pseudo-physical default probability, the average default risk premium is 56 percent in 1-year CDS during the normal period, it gradually increases with time horizon and eventually reaches 426 percent for the 10-year maturity.

Exhibit 5 gives a visual exhibition of the evolution of the whole term structure of the default risk premiums in Argentine debt during our sample period. The graph shows that the default risk premium steadily increased from January 1999 to September 2000, and there was a big jump in default risk premium around October 2000 when the market first realized a serious possibility of default. After October 2000, the default risk premium declined as the market expected default probability continued to rise dramatically. It is also obvious that the default risk premium is upward-sloped during the sample period.

In the case of Argentine default, there seems to be a turning point in the sense that when default probability is low, default risk premium increases with rising default probability. But when default probability reaches a certain point, default risk premium instead declines when default probability continues to ascend. These results indicate that when the default probability is moderate, the CDS buyer is willing to pay a relatively high premium to hedge for the default, and the default risk premium plays a primary role in the valuation of a CDS contract. On the other hand, when the pseudo-physical default probability rises above a certain level, the actual default probability becomes relatively more important.

We also implemented a GMM procedure to test the statistical significance of the difference between the two default probabilities and find that the risk-neutral default probability is significantly higher than its pseudo-physical counterpart during the sample period for all time-horizons.

We have shown that default risk premium varies from period to period. A natural question to ask is: What drives the variation in the default risk premium in Argentine debt over time? Bakshi Madan, and Zhang [2001] shows that the difference between these two default probabilities depends on the risk aversion of market participants, the magnitude of the pseudo-physical default probability, and the market expectation of the recovery rate under physical measure. Empirically, these factors may be related to changes in U.S. business cycles and credit conditions and the overall strength of the Argentine economy. To analyze causes of changes in default risk premum, we run the following regression:

$$\Delta DP(t) = \alpha_0 + \alpha_1 T10(t) + \alpha_2 Term(t) + \alpha_3 Credit(t) + \alpha_4 Merv(t) + \in(t) \qquad (21)$$

where $\Delta DP(t)$ is the wedge between the risk-neutral and the pseudo-physical default probabilities, **T-10** is the 10-year Treasury yield, **Term** is the spread between the 6-month and 10-year Treasury yields, **Credit** is the U.S. credit spread defined as the difference between the AAA and BBB corporate Merrill Lynch bond indexes, and **Merv** is the BUSE Merval stock price index of Argentina.

Panel A of Exhibit 6 reports the regression results. First, we note that the 10-year U.S. Treasury yield is significantly negative in the regression, while at the same time, the term premium of the U.S. term structure is positive and significant. These two results are consistent with

EXHIBIT 5

Default Risk Premium in CDS Prices on Argentine Sovereign Debt (01/20/1999–12/05/2001)

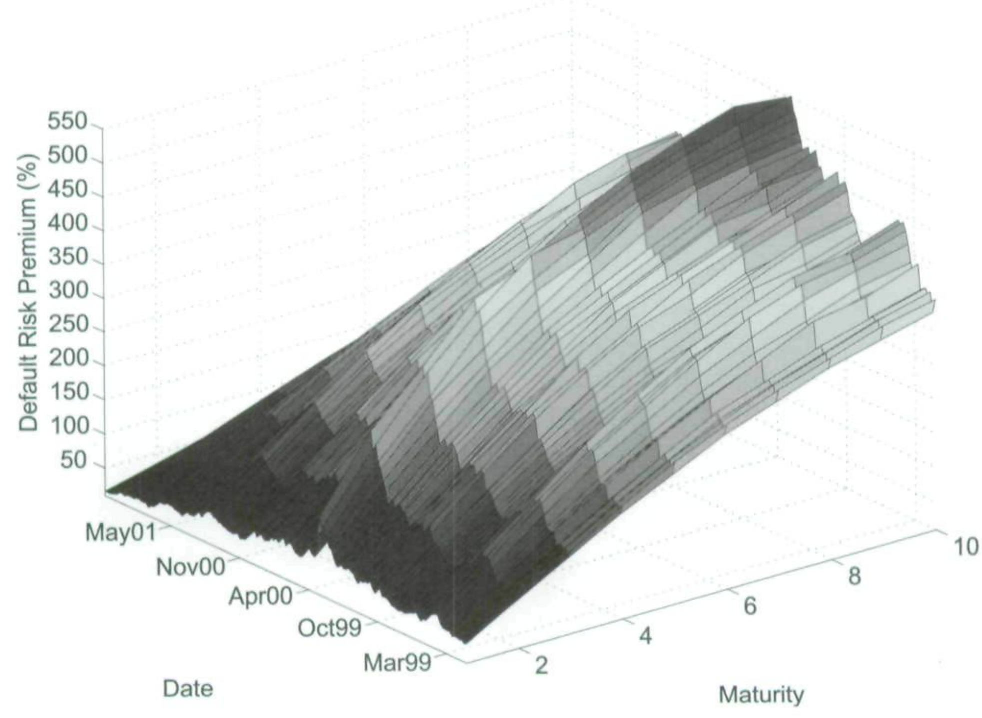


EXHIBIT 6

The Determinants of Default Risk Premium and Correlation Analysis

Panel A analyzes the variation of the difference between the risk-neutral and pseudo-physical 1-year default probabilities over time in the following regression: $\Delta DP(t) = \alpha_0 + \alpha_1 T10(t) + \alpha_2 Term(t) + \alpha_3 Credit(t) + \alpha_4 Merv(t) + \varepsilon(t)$, where **T-10** is the 10-year Treasury yield, **Term** is the spread between the 6-month and 10-year Treasury yields, **Credit** is the U.S. credit spread between the AAA and BBB corporate Merrill Lynch bond indices, and **Merv** is the BUSE Merval stock price index of Argentina. $\Delta DP(t)$ is the difference between the 1-year pseudo-physical and risk-neutral default probabilities. Nobs - Number of observations. Panel B presents results on correlation between the implied factors in our credit default swap model and relevant observable financial series in U.S. and Argentina. The three implied factors in the exhibit are the two U.S. term structure factors s_1, s_2, and the Argentina-specific factor **Z**. The financial series include **T-10**, **Term**, **Credit**, the spread between the 10-year US Treasury yield and the JP Morgan EMBI Index for Argentina (**Embi**), and **Merv**. The period is from 01/20/1999 to 12/05/2001. Data source: Bloomberg, Federal Reserve Board and JP Morgan Chase.

Panel A: The Determinants of Default Risk Premium

α_0	T-10	Term	Credit	Merv	R^2	Nobs
0.0494 (3.06)	-0.007 (-3.14)	0.014 (15.45)	0.021 (6.05)	-0.035 (-1.18)	0.86	148

Panel B: Correlation between Implied Factors and Selected Financial Series

Variable	s_1	s_2	T-10	Term	Credit	Z	Embi	Merv
s_1	1.000							
s_2	0.439	1.000						
T-10	0.565	0.898	1.000					
Term	-0.952	-0.423	-0.448	1.000				
Credit	-0.750	-0.588	-0.771	0.596	1.000			
Z	-0.852	-0.404	-0.532	0.820	0.739	1.000		
Embi	-0.808	-0.377	-0.505	0.778	0.730	0.984	1.000	
Merv	-0.128	-0.042	-0.041	0.163	0.046	0.181	0.214	1.000

the evidence that a large term premium of the term structure points to a steep yield curve, which is usually an indication of slow economic growth. Second, the U.S. credit premium is positive and significant. This result indicates that the Argentine credit condition is closely related to the U.S. domestic debt market and that the worsening of the U.S. credit conditions amplify the default risk premiums in Argentine sovereign debt. Third, the Argentine stock index return is negative, as expected, but it is insignificant. The variables on the right-hand side of the regression can explain about 86 percent of the time-series variation in default risk premiums. Taken together, the regression results suggest that the U.S. business cycle and credit condition and the overall domestic economic strength are important drivers of the default risk premium in Argentine sovereign debt.

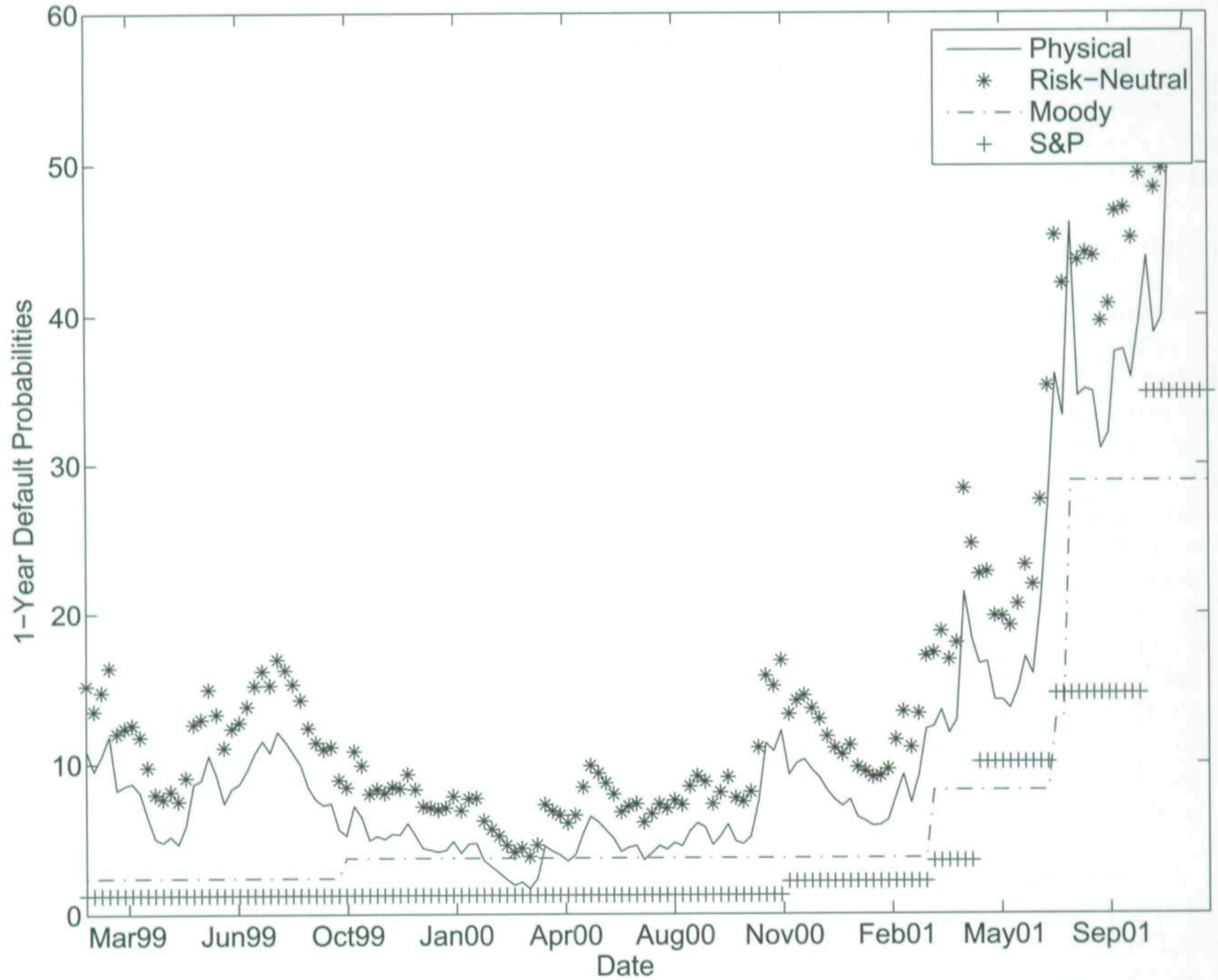
Rating-Implied Default Probabilities versus Market Expectations

In this sub-section, we examine whether the third-party rating agencies led or lagged the credit market in downgrading the Argentine debts. Because pseudo-physical default probability implied by our model and the historical default probability implied by credit rating are different entities, our emphasis is on the timing of rating agency actions instead of on comparing the magnitudes of the two probabilities. Here we focus on the 1-year implied default probability because the historical default rate data on credit ratings available from Moody's and S&P are of one-year horizons. We use default probabilities on corporate debts as an approximation for that on sovereign debts for two reasons. First, we don't have data on historical default probabilities on sovereign debts on sub-letter grades from either Moody's or S&P's. Second, comparing the default transition matrix on sovereign and corporate debts on letter grades provided by Moody's, it is obvious that the default transition matrix of corporate debts is more reliable than that of sovereign debts because the relation between default probabilities and ratings is not even monotone for sovereign debts.[9]

Exhibit 7 plots the 1-year pseudo-physical and risk-neutral default probabilities, as well as the default

EXHIBIT 7

1-Year Pseudo-Physical and Risk-Neutral Default Probabilities of Argentine Sovereign Debt



probabilities implicit in the ratings assigned to the Argentine sovereign debts by Moody's and Standard and Poor's over the entire sample period. One obvious feature of the graph is that the 1-year pseudo-physical and risk-neutral default probabilities are highly correlated, where the risk-neutral default probability moves closely together with its pseudo-physical counterpart. Exhibit 7 also shows that the 1-year risk-neutral default probability is always substantially higher than its pseudo-physical counterpart during the sample period. During most of the normal period from January 1999 to March 2001, the 1-year pseudo-physical default probability stayed below the 10% level, with lows around 2%. Around March 2001, however, the 1-year pseudo-physical default probability jumped significantly and further reached the magnitude of 30 to 40% around July 2001, and it eventually touched the level around 50% at the end of the sample period.

We now turn to assess whether rating agencies led or lagged the market in downgrading Argentine debts. We use actions by Moody's and S&P on the 10-year 11 percent fixed coupon Eurobond maturing on October 9, 2006, as the benchmark. Due to the lack of historical data on finer ratings in the broad rating class of Caa by Moody's (or CCC by S&P), we only examine actions by the two rating agencies up to the time when the debt was downgraded to the broad class of Caa (or CCC).

Exhibit 7 shows that before October 1999, Moody's assigned a rating of Ba3 to the Argentine debt. After seven months, Moody's downgraded Argentine debt from Ba3 to B1 on 10/06/1999. Moody's maintained the B1 rating on the bond until 03/28/2001, when it downgraded the bond to B2, about five months after the market-implied default probability jumped. Moody's then downgraded the bond again from B2 to B3 on 07/13/2001, and from B3 to Caa1 on 07/26/2001; both were behind the market reactions. Exhibit 7 shows that, compared to Moody's, all the downgrades made by S&P were way behind the market moves.

SPECIFICATION ANALYSIS

We show in earlier sections that the first extracted factor of the U.S. term structure is highly correlated with the negative slope of the Treasury curve and the second extracted term structure variable has a high correlation with the 10-year Treasury yield. In this subsection, we examine other correlations between the three extracted factors and a group of U.S. and Argentine macroeconomic and financial variables. Due to the high frequency nature of our CDS data, we focus on the financial variables that are available on a weekly basis and exclude those macroeconomic series that are available only on quarterly or annual frequencies, such as the debt/GDP ratio, the foreign exchange reserves, and other low frequency variables.

Panel B of Exhibit 6 displays the correlation between the three implied state variables and the following economic and financial variables: the 10-year U.S. Treasury yield, the term premium between the 6-month and the 10-year U.S. Treasury yields, the credit spread between the AAA and BBB Merrill Lynch U.S. corporate bond indexes, the spread between the 10-year U.S. Treasury yield and the JP Morgan EMBI Index for Argentina (Embi), and the weekly return on the Argentine stock index Merval (Merv).

Besides the high correlation results between the two U.S. term structure factors and the (negative) Treasury slope and level, Exhibit 6 shows that the implied Argentina-specific distress factor is highly correlated with the spread between the 10-year U.S. Treasury yield and the JP Morgan EMBI index with a correlation coefficient of 0.984. On the other hand, there is not much correlation between the Argentina-specific distress factor and the weekly Argentine stock index return. In fact, the correlation between the extracted third factor and the Argentine stock index return even comes up with a wrong sign. This result suggests that our extracted Argentina-specific distress factor captures default risk in Argentine sovereign debt pretty well.

Finally, some other strong correlations among the implied state variables and the selected economic variables are also noteworthy. For example, the first U.S. term structure factor is strongly and negatively correlated with the U.S. credit spread, the Argentina-specific distress factor, and the Argentina EMBI spread. Also, the implied Argentina-specific factor is strongly positively correlated with the U.S. term premium. Together, these results suggest that, at least for our sample period, when the U.S. term structure of interest rate is steep, credit conditions in both the U.S. debt market and the Argentine sovereign bond tend to worsen.

CONCLUSION

The fast emergence of the CDS market together with actual default events in the last several years provides us an excellent platform to explore the expectations of the credit market on the prospect of an eventual default. While the literature has seen a growing list of articles on CDSs, there are few empirical studies on the expectations of the credit market before a major default. Moreover, little is known about the term structure and time-series behavior of default risk premiums implicit in CDS prices. In this article, we examine these issues by developing a flexible CDS model and implementing the model to a unique data set of CDSs on Argentine sovereign debt that is rich in both cross-sectional and time-series observations.

Our empirical results show that both pseudo-physical and risk-neutral default probabilities increase dramatically from the start of our sample period to the months before the eventual Argentine default. We find that default risk premiums implicit in CDS prices are substantial. When default probability rose from a moderate level, default risk premium first increased and then declined. This result seems to suggest that when the default probability is low, CDS buyers are willing to pay a relatively high premium to hedge default. However, when the pseudo-physical default probability rises above a certain level, actual default probability dominates the pricing of the CDS contract and the effect of default risk premium becomes less important.

We also find that the term structures of the pseudo-physical and risk-neutral default probabilities may have different shapes. The pseudo-physical default probability is downward-sloped over the entire sample period, while the risk-neutral default probability initially has an upward-sloping curve during the majority of the sample period and the curve turns downward-sloping when default risk becomes extremely high toward the end of the sample period. The term structure of default risk premiums is upward-sloping throughout the sample period. Regression analysis suggests the default risk premium was affected by changes in the U.S. business cycle and credit conditions and the overall strength of Argentine economy. Comparing with the market expected default probabilities, we find that both Moody's and Standard and Poor's lagged the market in downgrading the debt.

APPENDIX

Proof of Proposition 1

From Equations (5), (9), and (11), characteristic function in (11) can be written as

$$\begin{aligned}\Phi(t,\tau;\phi) &= e - (\alpha_r + \Lambda_0 + \varphi_0 - \Lambda_{s_1}\bar{s}_1 \\ &\quad - \Lambda_{s_2}\bar{s}_2 - \varphi_{s_2}\bar{s}_2)\tau + i\varphi\,(\Lambda_0 - \Lambda_{s_1}\bar{s}_1 - \Lambda_{s_2}\bar{s}_2) \\ &\quad \times \Phi_1(t,\tau;\phi) \times \Phi_2(t,\tau;\phi) \times \Phi_3(t,\tau;\phi)\end{aligned}$$

where

$$\Phi_1(t,\tau;\phi) = E_t^Q\left[e^{-(1+\Lambda_{s_1}+\varphi_{s_1})\int_t^{t+\tau} s_1(u)du + i\phi\Lambda_{s_1}s_1(T)}\right]$$

$$\Phi_2(t,\tau;\phi) = E_t^Q\left[e^{-(1+\Lambda_{s_2}+\varphi_{s_2})\int_t^{t+\tau} s_2(u)du + i\phi\Lambda_{s_2}s_2(T)}\right]$$

$$\Phi_3(t,\tau;\phi) = E_t^Q\left[e^{-\int_t^{t+\tau} Z(u)du + i\phi Z(T)}\right]$$

Functions $\Phi_1(t,\tau;\phi)$, $\Phi_2(t,\tau;\phi)$ and $\Phi_3(t,\tau;\phi)$ can be solved in closed-form which is exponential-affine.

Probabilities of Survival of Underlying Reference

The probability of survival of the underlying reference from time t to $t+\tau$, under the risk neutral measure, is

$$\begin{aligned}G(t,\tau) &= E_t^Q[e^{-\int_t^{t+\tau} h_1(u)du}] \\ &= e^{A_G(\tau) - B_G(\tau)s_1(t) - C_G(\tau)s_2(t) - D_G(t,\tau;\phi)Z(t)}\end{aligned} \tag{22}$$

where $A_G()$, $B_G()$, $C_G()$ and $D_G()$ are closed-form functions. The pseudo-physical probability of survival can be obtained likewise.

Expressions $f_z(\hat{Z}_t|\hat{Z}_{t-1})$ and $|J_t^c|$ in the Log-Likelihood Function

In the log-likelihood function in Equation (20), $f_z(\hat{Z}_t|\hat{Z}_{t-1})$ is the conditional density $\hat{Z}_t$ given $\hat{Z}_{t-1}$, and J_t^C is the Jacobian of variable transformation,

$$J_t^C = \frac{\mathcal{N}^C}{\left[\int_t^{t+5}\Phi(t,u;\phi=0)du\right]^2}$$

with

$$\begin{aligned}\mathcal{N}^C &= \left[\int_t^{t+5}\gamma_1 \frac{\partial\left(\frac{1}{i}\frac{(\partial\Phi(t,u;\phi)}{\partial\Phi}|_{\phi=0}\right)}{\partial Z_t}du\right] \times \left[\int_t^{t+5}\Phi(t,u;0)du\right] \\ &\quad + \left[\int_t^{t+5}\gamma_1\left(\frac{1}{i}\frac{\Phi(t,u;\phi)}{\partial\Phi}|_{\phi=0}\right)du\right] \times \left[\int_t^{t+5}\Phi(t,u;0)D(t,u;0)du\right]\end{aligned}$$

and

$$\begin{aligned}\frac{\partial\left(\frac{1}{i}\frac{(\partial\Phi(t,u;\phi)}{\partial\Phi}|_{\phi=0}\right)}{\partial Z_t} &= -\Phi(t,u;0)D(t,u;0)\left[\frac{1}{i}\frac{\partial A}{\partial\phi}|_{\phi=0}\right. \\ &\quad - \frac{1}{i}\frac{\partial B}{\partial\phi}|_{\phi=0}s_1 - \frac{1}{i}\frac{\partial C}{\partial\phi}|_{\phi=0}s_2 \\ &\quad \left. - \frac{1}{i}\frac{\partial D}{\partial\phi}|_{\phi=0}Z\right] - \Phi(t,u;0)\frac{1}{i}\frac{\partial D}{\partial\phi}|_{\phi=0}\end{aligned}$$

In the implementation of the QML estimation, we substitute the exact transition density $f_z(Z_t|Z_{t-1})$ with a normal density: $Z(t)\,|\,Z(t-1) \sim N(\mu_z(t), Q_z(t))$, where $\mu_z(t) = \theta_z(1-e^{-\kappa_z\Delta t}) + e^{-\kappa_z\Delta t}Z(t-1)$, and $Q_z(t) = \sigma_z^2 k_z^{-1}(1-e^{-\kappa_z\Delta t})[\frac{\theta_z}{2}(1-e^{-\kappa_z\Delta t}) + e^{-\kappa_z\Delta t}Z(t-1)]$.

ENDNOTES

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[1]The τ-year conditional cumulative pseudo-physical default probability is calculated as, $1-E^P[e - \int_t^{t+\tau} h(s)ds]$, where $h(s)$ is the time-s risk-neutral instantaneous hazard rate, and the expectation is taken under the physical measure P (see the Appendix for details).

[2]A partial list of recent theoretical contributions include Das and Sundaram [2000], Duffie [1999a], Hull and White [2000a], Hull and White [2000b] and Jarrow and Yildirim [2002].

[3]Though easily relaxed, we are making the implicit assumption that there is no default by the CDS buyer. We also assume default by CDS seller on its other debts is exogenous to CDS seller.

[4]For a review of the recent history of the political and economic events in Argentina, see Mussa [2002] and Pando [2002].

[5]We also estimated all parameters, including that of the default-free interest rates and the default swaps simultaneously in a single step. But due to the large number of parameters and the highly non-linear nature of the formulas, the estimation is not as stable as the two-step procedure.

[6]There are two reasons for this choice. First, there are concerns about whether Treasury yields should still be viewed as the benchmark due to the dwindling trading in Treasury securities since 1998. Furthermore, Treasury rates also differ from the "true" risk-free rates because of the repo effects, liquidity differences, and tax shields (see Collin-Dufresne and Solnik [2001] and Duffie and Singleton [1997]). For this reason, U.S.-dollar LIBOR-quality swap yields have gained popularity among both practitioners and academia as the new benchmark of the risk-free reference rates. Second, interest rate swaps data are widely available at a range of constant times to maturity, which makes estimation process less than complicated. Although interest rate swaps are defaultable in theory, the effects of the counter-party default risk of swap contracts are believed to be minimal (see Duffie, Pedersen, and Singleton [2003]).

[7]Estimation details can be obtained from authors if interested.

[8]Alternatively, one can also include the log of the unconditional log likelihood to construct the exact log-likelihood function. Given that the conditional MLE and the exact MLE have the same large sample distributions and that the conditional MLE provides consistent estimates under some circumstances while the exact MLE does not, we chose to use the conditional MLE method (see Hamilton [1994]).

[9]In our study, Moody's one-year default rates are based on period from 1981 to 2003 and those of S&P are from 1983 to 2003.

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