Ambiguity Shifts and the 2007–2008 Financial Crisis*

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Abstract

I analyze the effects of model misspecification on default swap spreads and equity prices for firms that are informationally opaque to the investors. The agents in the economy are misspecification-averse and thus assign higher probabilities to lower utility states. This leads to higher CDS rates, lower equity prices and lower expected times to default. Estimating the model using data on financial institutions, I find that the sudden increase in credit spreads in the summer of 2007 can be partially explained by agents' mistrust of the signals observed in the market. The bailout of Bear Stearns in March 2008 and the liquidation of Lehman Brothers in September 2008 further exacerbated the agents' doubts about signal quality and introduced mistrust about the agents' pricing models, accounting for the further increases in credit spreads after these events.

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1 Introduction

When making consumption decisions, an investor faces uncertainty about both the relevant underlying state and the data-generating process governing the evolution of the state. While uncertainty about the state is risk that the investor understands and can model, uncertainty about the data-generating process represents agents' pessimism about their ability to identify the correct model. This paper argues that prices of credit securities are sensitive to the investors' preferences toward model uncertainty and that the implied time-variation in the level of model uncertainty is a source of variation in credit spreads that explains the asymmetric response of credit spreads to upturns and downturns in the economy.

I analyze the effects of model misspecification on default swap spreads in secondary markets for the corporate debt of firms that are not perfectly transparent to the investors. The agents in the economy are misspecification-averse and thus mistrust the statistical model of the fundamental value of assets of firms and of the firms' observed earnings process. This mistrust reflects the fact that, when it is difficult for investors to observe firms' assets directly, they are forced to rely on imprecise accounting information. In this situation, investors must draw inference from accounting data and other publicly available information. Investors realize that, although they may be able to pick a model of the fundamental asset value and the accounting signals to best fit the historical data, this may not be the true data-generating model. Under the assumption that investors are misspecification-averse, I derive the asset prices in the economy, explicitly accounting for the implications of imperfect information and model misspecification.

I show several significant implications of model misspecification for the level and variation in the term structure of secondary market default swap spreads. Compared to a model with perfect information, model uncertainty increases the level of the yield curve and the default-swap spread curve. Intuitively, in the presence of model misspecification, investors must be compensated for the risk associated with choosing the "wrong" model to describe the evolution of the underlying state. Notice that, as shown in Duffie and Lando [2001], introducing imperfect information to a standard Black and Cox [1976] model has the additional benefit of being able to explain high credit spreads at short maturities.

Next, I show that model uncertainty exacerbates the imperfect information problem faced by the representative investors in the secondary asset markets. In filtering information about the underlying state from the imperfect signals, agents must take into account uncertainty about both the model governing the evolution of the underlying state and the signals about the underlying state. The misspecification-averse agent assigns higher probabilities to lower utility states. Further, how much these probabilities are higher than under the reference model depends on the current conditional probability vector under the reference model.

Model misspecification also impacts the joint probability distribution of the next period's signals and states. In particular, while in states of the economy when no firm defaults, the misspecification averse agent perceives the probability of transitioning to a default state to be higher than under the reference model. Thus, the expected time to default of each firm decreases, increasing default-swap spreads. Further, the misspecification averse agent also perceives the transition probability matrix associated with the underlying state to be time varying. The time-variation in the transition probability matrix induces additional time-variation in the expected time to default of each firm and, thus, in default swap spreads.

In this paper, I argue that the increases in CDS spreads observed during the 2007–2008 crisis were due to increases in investors' doubts about the validity of their pricing models and the quality of the signals available to market participants. On August 9, 2007, France's largest bank BNP Paribas announced that it was having difficulties because two of its off-balance-sheet funds had loaded up on securities based on American subprime mortgages. But Paribas was not alone in its troubles: a month before, the German bank IKB announced similar difficulties, and the Paribas announcement was followed the next day by Northern Rock's revelation that it had only had enough reserve cash to last until the end of the month. These and other similar announcements lead to a freeze of the credit markets as banks lost faith in each other's balance sheets. The situation was particularly surprising considering the market conditions shortly before the crisis began. At the beginning of 2007, financial markets were liquidity-unconstrained and credit spreads were at historical lows. Even as late as May 2007, it would have been hard to predict the magnitude of the response that the losses on subprime mortgages had generated. Compared to the total value of financial instruments traded worldwide, the subprime losses were relatively small: even the worst-case estimates put them at around USD 250 billion. [1] Further, for investors familiar with the instruments, the losses were not unexpected. By definition, the subprime mortgages were part of the riskiest segment of the mortgage market, so it was hardly surprising some borrowers would default on the loans. Yet, despite their predictability, the defaults had precipitated the current liquidity crisis that spread between the credit markets.

Using observations of the CDS spreads on financial institutions, I estimate the degree of misspecification aversion of the investors in the secondary debt markets. To evaluate the changes in investors' aversion to misspecification during the crisis, I estimate the misspecification aversion coefficient using three sub-periods– before the start of the crisis in July 2007, from the start of the crisis to the bailout of Bear Stearns in March 2008, and from the bailout of Bear Stearns to the liquidation of Lehman Brothers in September 2008 – and find that the three estimates are not statistically significantly different, implying that investors' misspecification preferences did not change during the crisis. In terms of the model, this implies that the observed changes in credit spreads during the financial crisis were due not to changing misspecification attitudes on the part of the investors but rather due to an increase in the amount of misspecification in the economy.

Since the investors' aversion to misspecification did not change significantly during the crisis, I use the pre-crisis estimate of the misspecification aversion coefficient to compute the model-implied time series of CDS spreads, equity prices and the quantity of misspecification. Measuring the amount of misspecification using the expected log likelihood ratio (or relative entropy) between the reference and the worst-case models, I find that the amount of misspecification did in fact increase during the financial crisis. Further, the way that total entropy is decomposed into the contribution from misspecification of the distribution of the future signals and state and the contribution from misspecification of the conditional distribution over the current state

[1] Source: Caballero and Krishnamurthy [2008b]

changed during the crisis. More specifically, the initial BNP Paribas announcement in August 2007 lead to an increase in the relative entropy due to current period conditional probability misspecification. Intuitively, the BNP Paribas announcement and the subsequent Northern Rock revelations lead to an increase in ambiguity about the quality of the signals available to market participants. The bailout of Bear Stearns and the bailout of Lehman Brothers, on the other hand, lead to an increase in both components of entropy. That is, the effective default of these two institutions increased not only investors' doubts about the quality of the surviving financial institutions (increasing entropy due to misspecification of distribution of future signals and states), but also investors' doubts about the quality of the signals in the market. Entropy stabilizes toward the end of the crisis but at a higher level than before the start of the crisis.

To evaluate the quality of the fit of the model, I compare the model-implied CDS spreads and equity prices to the corresponding quantities observed in the data. The model-implied CDS spreads match both the levels and the changes in CDS spreads observed during the crisis, although the performance of the model deteriorates after the liquidation of Lehman Brothers. Further, although the model does not match the overall levels of equity prices and, in fact, is not geared to do so, it does match the changes in equity prices observed during the crisis.

The rest of the paper is organized as follows. I review the related literature in Section 2. I describe the model considered in the paper in Section 3. The results of the estimation of the model are presented in Section 4. Section 5 concludes. Technical details are relegated to the appendix.

2 Literature Review

A rapidly growing literature studies the behavior of asset prices in the presence of ambiguity in dynamic economies. A substantial part of this literature considers investor ambiguity about the data-generating model. Anderson et al. [2003] derive the pricing semigroups associated with robust perturbations of the true state probability law. Trojani and Vanini [2002] use their framework to address the equity premium and the interest rate puzzles, while Leippold et al.

[2008] consider also the excess volatility puzzle. Gagliardini et al. [2009] study the term structure implications of adding ambiguity to a production economy. This setting has also been used to study the portfolio behavior of ambiguity-averse investors and the implications for the options markets (see e.g. Trojani and Vanini [2004] and Liu et al. [2005]).

The second strand in the literature, however, assumes that, although the agents in the economy know the "true" data-generating model, they face uncertainty about the quality of the observed signal about an unobservable underlying. Chen and Epstein [2002] study the equity premium and the interest rate puzzles in this set-up, and Epstein and Schneider [2008] consider the implications for the excess volatility puzzle. The portfolio allocation implications of this setting have also been studied extensively in e.g. Uppal and Wang [2003] and Epstein and Miao [2003].

However, none of these papers study the relationship between ambiguity aversion and the term structure of credit spreads. Following Hansen and Sargent [2005], Hansen and Sargent [2007], I introduce model misspecification by considering martingale distortions to the reference model probability law. As Hansen and Sargent [2007] show, the martingale distortion can be factored into distortions of the conditional distribution of the underlying state (signal qulaity) and the evolution law of the hidden state (asset value dynamics). I assume that the representative investor in the secondary debt market has max-min preferences over consumption paths under possible models.

This paper is also related to the literature on preference-based explanations for credit spreads. Chen [2010] studies two puzzles about corporate debt: the credit spread puzzle – why yield spreads between corporate bonds and treasuries are high and volatile – and the under-leverage puzzle – why firms use debt conservatively despite seemingly large tax benefits and low costs of financial distress. The paper argues that both of these puzzles can be explained by two observations: defaults are more highly concentrated during bad times, when marginal utility is high, and the losses associated with default are higher during such times. Thus, investors demand high risk premia for holding defaultable claims, including corporate bonds and levered firms.

Using similar intuition, Chen et al. [2009] argue that the credit spread puzzle can be explained by the covariation between default rates and market Sharpe ratios. That is, investors must be compensated more for holding credit risk securities because default rates (and, hence, expected losses from default) increase at the same time as market returns are more uncertain. More specifically, the authors investigate the credit spread implications of the Campbell and Cochrane [1999] pricing kernel calibrated to equity returns and aggregate consumption data. Identifying the historical surplus–consumption ratio from aggregate consumption data, the paper finds that the implied level and time-variation of spreads match historical levels well.

3 Model

In this section, I present the economy considered in this paper. I begin by describing the reference model used for pricing credit securities and then proceed to the misspecification problem faced by the representative agent in the economy.

3.1 Reference model

As the reference model, I consider a modified version of the Black and Cox [1976] economy. Consider a (sector of the) economy consisting of I firms, indexed by $i = 1, \ldots, I$ and denote by $A_{it} = e^{a_{it}}$ the fundamental value of the assets of firm i at date $t = 1, 2, \ldots$. To fix ideas, assume that there are $n_y = 12$ data periods in a year, so that each period corresponds to a month. I assume that the log-asset value of each firm can be decomposed into the sum of two components:

$$a_{it} = z_{it} + \rho_i z_{ct}, \tag{3.1}$$

where z_{it} is an idiosyncratic shock to the asset value of firm i, z_{ct} is an aggregate shock to the asset values of all the firms in the sector and ρ_i is the loading of firm i on the aggregate component. Denote by $z_t = [z_{1t}, \ldots, z_{It}, z_{ct}]'$ the vector of the components of asset values at date t. I assume that the vector z_t evolves according to an N-state Markov chain, with possible values $\xi_1, \ldots, \xi_N$ and the transition probability matrix Λ defined as:

$$\{\Lambda\}_{jk} \equiv \lambda_{jk} = \mathbb{P}\left(z_{t+1} = \xi_k \middle| z_t = \xi_j \right). \tag{3.2}$$

There are two types of agents in the economy: managers and investors. All the day-to-day operations of the firm are delegated to the respective manager. I assume that there are no agency problems between a firm's managers and the equity holders of the firm, so that the managers act in the best interest of the equity holders. Further, similarly to Duffie and Lando [2001], I assume that the managers are better informed about the firm they manage than the participants in the public markets and, in particular, that the managers of the firm observe perfectly the evolution of the fundamental value of the firm's assets. To prevent information spill-over, I assume that managers are precluded from trading in the public assets markets.

In this paper, I abstract from modeling the operational decisions of the firm managers and, in particular, from modeling the optimal capital structure, dividend payment and default decisions faced by the managers. As in Leland [1994] and Duffie and Lando [2001], I assume that each firm i issues perpetual debt with face value D_i. This debt is serviced by a constant coupon rate C_i. While firm i is in operation, it generates a constant fraction δ_i of assets as cash-flows which accrue, minus the coupon payments, as equity in the firm.

The managers decide on behalf of the equity holders when to default. As in Black and Cox [1976], I abstract from modeling the liquidation decision faced by the managers and assume instead that a firm defaults automatically whenever the fundamental value of the firm's assets reaches the lowest possible value implied by the Markov chain $\{z_t\}_{t=1}^{+\infty}$. In particular, denote by ξ_{ji} the i^{th} element of the asset values vector in state j, ξ_j. Let $i^* = argmin_{j=1,\ldots,N}\xi_{ji} + \rho_i \xi_{jc}$ be the state index at which firm i achieves its lowest possible value and by $a_{B_i} = \xi_{i^*}$ the corresponding state. Then the (stochastic) default date τ_i of firm i is the first hitting time of the state a_{B_i}: $\tau_i = \inf\{t : z_t = a_{B_i}\}$. In economic terms, the exogenous default rule can be interpreted as a debt covenant. The firm is liquidated at the present value of the discontinued cash flows, with the proceeds distributed among the firm's primary debt holders and the equity holders receiving 0. For simplicity, I assume that each firm has a single default state and that firms do not default simultaneously. Notice that, since managers observe perfectly the asset value evolution of the firm under their management, there is no uncertainty about the firm being liquidated upon hitting its default boundary. Finally, denote by $a_B = \bigcup_{i=1}^{I} a_{B_i}$ the union of the default states of all the firms and by a_B^c its complement, which is the set of states where none of the firms default.

Consider now the participants in the public markets. Similarly to Duffie and Lando [2001], I assume that the representative investor does not observe the true evolution of asset values in the sector and receives instead imperfect, unbiased signals about the fundamental value of the assets of each firm, $\hat{A}_{it} = e^{y_{it}}$, and the aggregate component of asset values in the sector, $\hat{A}_{ct} = e^{y_{ct}}$. More specifically, assume that $y_{it} = a_{it} + u_{it}$ and $y_{ct} = z_{ct} + u_{ct}$ where the signal errors $u_t = [u_{1t}, \ldots, u_{It}, u_{ct}]'$ are serially uncorrelated and normally distributed, independent of the true realization of z_t: $u_t \sim N(\overline{u}, \Sigma_u)$. Here, $\overline{u}$ is the mean signal error and Σ_u^{-1} the signal quality. At each date t, the representative agent also observe whether any of the default states have been reached and any of the firms have been liquidated. Thus, the information set of the representative agent at date t is:

$$\mathcal{G}_t = \sigma\left\{y_s,\ \mathbf{1}_{z_s \in a_B^c}:\ s = 1, \ldots, t\right\},$$

where $y_t = [y_{1t}, \ldots, y_{It}, y_{ct}]'$ is the full signal vector at date t.

Denote by p_{jt} the probability, conditional on the date t information set of the representative investors, of the vector z being in state j at date t:

$$p_{jt} = \mathbb{P}\left(z_t = \xi_j \middle| \mathcal{G}_t\right).$$

For mathematical reasons, it is easier to formulate the updating rule in terms of unnormalized probabilities $\vec{\pi}_t$, which are related to the proper probabilities by:

$$p_{jt} = \frac{\pi_{jt}}{\sum_{j=1}^{N} \pi_{jt}}, \qquad j = 1, \ldots, N.$$

Let $\pi_{j0} = \mathbb{P}(z_1 = \xi_j)$ be the prior probability. Then the following result holds.

Lemma 3.1. *(Wonham Filter)*

Assume that the transition probability matrix, Λ, and the prior distribution π_{j0} are known. Then the date 1 update to the unnormalized probability vector is given by:

$$\pi_{j1} = \mathbf{1}_{\xi_j \notin a_B} \pi_{j0} f(y_1 - \xi_j), \qquad j = 1, \ldots, N, \tag{3.3}$$

where $f(\cdot)$ is the density function of the observation errors, u. For $t > 1$, the "predict" step of the update to the unnormalized probability vector is given by:

$$\tilde{\pi}_{t+1} = diag(\mathbf{f}(y_{t+1}))\Lambda' \vec{\pi}_t, \tag{3.4}$$

where $\mathbf{f}(y) = [f(y - \xi_1), \ldots, f(y - \xi_N)]'$ and $diag(\cdot)$ creates a diagonal matrix from the vector $\cdot$. Then, conditional on no firm defaulting in period $t+1$, the updated unnormalized probability vector is given by:

$$\vec{\pi}_{t+1} = diag(\mathbf{1}_{a_B^c})\tilde{\pi}_{t+1},$$

where $\mathbf{1}_{a_B^c} = \left[\mathbf{1}_{\xi_1 \in a_B^c}, \ldots, \mathbf{1}_{\xi_N \in a_B^c}\right]'$.

Proof. See e.g. Frey and Schmidt [2009]. □

Finally, consider the utility of the representative agent. I assume the representative agent is risk-neutral and, thus, holds all the claims to the firm's asset value. Thus, the date t expected present value of the utility of the representative agent is given by:

$$J_t = \mathbb{E}\left[\sum_{s=0}^{+\infty} \beta^s \sum_{i=1}^{I} \delta_i A_{i,t+s} \middle| \mathcal{G}_t\right], \tag{3.5}$$

where β is the subjective discount factor. For the discussion below, it is useful to represent the expected present value of utility in recursive form:

$$J_t = \mathbb{E}\left[\sum_{i=1}^{I} \delta_i A_{it} + \beta \mathbf{1}_{z_{t+1} \in a_B^c} J_{t+1} + \beta \sum_{i=1}^{I} \mathbf{1}_{z_{t+1} = a_{B_i}} J^{\tau_i} \middle| \mathcal{G}_t \right], \tag{3.6}$$

where J^{τ_i} is the value function of the representative agent in case firm i defaults.

3.2 Asset prices

In this paper, I consider three types of claims to the assets of the firm: a claim to the firm's equity, a zero-coupon, risky bond and a credit default swap (CDS) written on the bond. Recall that equity in firm i accrues as a constant fraction δ_i of the fundamental asset value and that the equity holders receive 0 in case of the firm being liquidated. Thus, the date t price of a claim to equity of firm i, V_{it}, is given by:

$$V_{it} = \mathbb{E}\left[\sum_{s=0}^{\tau_i} \beta^s (\delta_i A_{i,t+s} - C_i) \middle| \mathcal{G}_t \right],$$

where τ_i is the (stochastic) default date of the firm i. Notice that the equity price satisfies the Euler equation:

$$V_{it} = \mathbb{E}\left[\delta_i A_{it} - C_i + \beta \mathbf{1}_{\tau_i > t+1} V_{i,t+1} \middle| \mathcal{G}_t \right]. \tag{3.7}$$

Consider now the default-swaps written on the primary debt of firm i. With a given maturity T, a default-swap is an exchange of an annuity stream at a constant coupon rate until maturity or default, whichever is first, in return for a payment of X at default, if default is before T, where X is the difference between the face value and the recovery value on the stipulated underlying bond. A default swap can thus be thought of as a default insurance contract for bond holders that expires at a given date T, and makes up the difference between face and recovery values in the event of default.

I assume, as typical in practice, that the default-swap annuity payments are made semiannually, and that the default swaps maturity date T is a coupon date. As in Duffie and Lando [2001], I take the underlying bond for the default swap on firm i to be the consol bond issued by firm i. Recall that, in case of default, the debt holders receive the present value of the discontinued cash flows. Thus, the payment X_i per unit of primary debt if firm i defaults before the swap maturity date T is given by:

$$X_i = 1 - \frac{\delta_i (I - \beta\Lambda)^{-1}_{i*} e^{\xi}}{(1-\beta) D_i}.$$

The at-market default-swap spread is the annualized coupon rate $c_i(t,T)$ that makes the default swap sell at date t for a market value of 0. Thus, with $T = t + 6n$ for a given positive integer n,[2] the CDS spread is given by:

$$c_i(t,T) = \frac{2X_i \mathbb{E}\left[\beta^{\tau_i - t} \mathbf{1}_{\tau_i < T} \middle| \mathcal{G}_t\right]}{\sum_{s=1}^{n} \beta^{6s} \mathbb{E}\left[\mathbf{1}_{\tau_i \geq t+6s} \middle| \mathcal{G}_t\right]}. \tag{3.8}$$

Default-swap spreads are a standard for price quotation and credit information in bond markets. In this setting, they have the additional virtue of providing implicitly the term structure of credit spreads for par floating-rate bonds of the same credit quality as the underlying consol bond, in terms of default time and recovery at default. Denote by $B_i(t,T)$ the date t price of a zero-coupon bond with maturity date T on the debt of firm i. The implicit discount curve is then given by:

$$\begin{aligned} B_i(t, t+6) &= \frac{1}{1 + c_i(t, t+6)} \\ B_i(t, t+6s) &= \frac{1 - \frac{c_i(t,t+6s)}{2} \sum_{j=1}^{s-1} B_i(t, t+6j)}{1 + c_i(t, t+6s)/2}. \end{aligned}$$

[2] Recall that there are 12 data periods in a year

3.3 Model misspecification

This paper studies asset prices in a setting where the representative agent makes decision rules robust to possible misspecifications of asset value and accounting signals models. In reality, the correct specification assumption of the reference model is overly restrictive. It implies that, even though the participants in the public markets only observe imperfect signals about the evolution of the fundamental asset value, they can still correctly identify the parametric model that governs the relevant dynamics. More realistically, I assume that the representative investor in the firm fears misspecification of the probability law generated by the model above and believes instead that the signals are related to the true asset value realizations by a family of likelihoods.

As in Hansen and Sargent [1995], Hansen et al. [1999], Tallarini Jr. [2000] and Anderson et al. [2003], I model preferences of the representative agent in the presence of model misspecification using the recursion:

$$J_t = -\theta \log \mathbb{E}\left[\exp\left[-\frac{U(z_t) + \mathcal{R}_t(\beta J_{t+1})}{\theta}\right]\middle| \mathcal{G}_t\right], \tag{3.9}$$

where:

$$\mathcal{R}_t(\beta J_t) \equiv -\theta \log \mathbb{E}\left[\exp\left(-\frac{\beta J_{t+1}}{\theta}\right)\middle| \mathcal{F}_t\right].$$

The risk-sensitive recursion (3.9) replaces the standard utility recursion (3.6), incorporating the representative agent's misspecification doubts in two steps.[3] First, the tilted continuation function $\mathcal{R}_t$ makes an additional risk adjustment to the continuation value function of the representative agent, accounting for misspecification fears about the fundamental asset value evolution dynamics. Second, the tilted expectations over the current period utility adjusts

[3]Notice that the recursion (3.9) can be rewritten as:

$$J_t = -\theta \log \mathbb{E}\left[\exp\left[-\frac{U(z_t) + \beta J_{t+1}}{\theta}\right]\middle| \mathcal{G}_t\right].$$

The recursion (3.9) allows for easier interpretation of the corresponding worst-case likelihood. Further, the more general form (3.9) allows, for the representative agent to have different attitudes toward misspecification of future signals and state and misspecification of the conditional distribution over the current state. For more details on this formulation, see Hansen and Sargent [2007].

for misspecification fears about the filtered probability distribution over the current state. As emphasized by Hansen and Sargent [1995], the log-exp specification of the recursion links risk-sensitive control theory and a more general recursive utility specification of Epstein and Zin [1989]. The degree of the representative agent's aversion to misspecification is quantified by θ^{-1}. When $\theta^{-1} = 0$, the risk-sensitive recursion (3.9) reverts to the usual utility recursion under Von Neumann-Morgenstern form of state additivity. For values of θ^{-1} greater than zero, the recursion (3.9) implies an increased aversion to risk *vis a vis* the Von Neumann-Morgenstern specification. Maenhout [2004] links the degree of misspecification to the value function itself, so that the agent becomes more misspecification-averse as the present value of her utility increases.

To understand better the recursion (3.9), consider the following static optimization problem:

$$\min_{m \geq 0;\ \mathbb{E}[m]=1} \mathbb{E}\left[mV\right] + \theta \mathbb{E}\left[m \log m\right]. \tag{3.10}$$

The random variable m is the likelihood ratio between the reference model and an alternative model. m implies a distorted expectation operator: $\tilde{\mathbb{E}}\left[V\right] = \mathbb{E}\left[mV\right]$. The optimization problem (3.10) then minimizes the expected value of the payoff V under alternative models but is penalized in utility terms for deviations from the reference model (parametrized by $m = 1$). The term $\mathbb{E}[m \log m]$ measures the discrepancy in relative entropy terms between the reference model and an alternative model. As noted in **?**, the relative entropy $\mathbb{E}\left[m \log m\right]$ is the expected log-likelihood between the reference and the misspecified models. Thus, the parameter θ can be interpreted as a penalization parameter for large deviations away from the reference model. The problem (3.10) can thus be interpreted as a robust way of alternating probability measures. The minimizing choice of m, the so-called worst-case model, is given by:

$$m^* = \frac{\exp\left(-\frac{1}{\theta}V\right)}{\mathbb{E}\left[\exp\left(-\frac{1}{\theta}V\right)\right]}$$

and the outcome of the minimization problem by:

$$-\theta \log \mathbb{E}\left[\exp\left(-\frac{1}{\theta}V\right)\right].$$

Thus, in choosing between alternative models, the representative agent tilts the probability toward bad (in terms of payoffs) states.

Turn now back to the recursion (3.9). Hansen and Sargent [2007] show that, corresponding to the tilted continuation function $\mathcal{R}_t$ is the worst-case likelihood ratio:

$$\phi_t(z_{t+1}, y_{t+1}) = \frac{\exp\left(-\frac{\beta J_{t+1}}{\theta}\right)}{\mathbb{E}\left[\exp\left(-\frac{\beta J_{t+1}}{\theta}\right)\Big| \mathcal{F}_t\right]}. \tag{3.11}$$

ϕ_t captures the difference between the evolution of future signals and states under the misspecified model and under the reference model. In particular, ϕ_t is the date t probability distortion to the joint distribution of next period's signals and state. Relative to the reference model distribution, ϕ_t tilts the joint distribution toward lower continuation value states, decreasing the expected future value of the continuation utility. Similarly, corresponding to the recursion (3.9), is the worst-case likelihood ratio:

$$\psi_t(z) = \frac{\exp\left(-\frac{U(z)+\mathcal{R}_t(\beta J_{t+1})}{\theta}\right)}{\mathbb{E}\left[\exp\left(-\frac{U(z)+\mathcal{R}_t(\beta J_{t+1})}{\theta}\right)\Big| \mathcal{G}_t\right]} \tag{3.12}$$

between the conditional distribution over the state at date t under the misspecfied and reference models. ψ_t tilts the conditional distribution toward lower utility states, decreasing the expected value of utility at the current date.

An alternative interpretation of the recursion (3.9) is in terms of the smooth robustness preferences of Klibanoff et al. [2005] and Klibanoff et al. [2009] and recursive preferences of Epstein and Zin [1989]. In the smooth ambiguity preference setting, the representative agent does not choose the "worst-case model" and instead assigns a preference ordering to the alternative models. In particular, let u be the agent's utility over realizations of consumption, μ index different models, f the agent's utility function over different models and π the belief vector over the different models. Then, an agent with the smooth robustness preferences evaluates consumption according to:

$$f^{-1}\left(\mathbb{E}_{\pi}\left[f\left(\mathbb{E}\left[u|\,\mu\right]\right)\right]\right).$$

Compare this to the recursion (3.9). Notice first that the tilted continuation utility, $\mathcal{R}(J_{t+1})$, corresponds to the continuation utility for an agent with Epstein and Zin [1989] preferences, so corresponding to the inner expectation conditional on a "model" in smooth utility preferences, the worst-case utility model evaluates future consumption using Epstein and Zin [1989] preferences conditional on the current state. The recursion (3.9) then uses an exponential utility function to rank continuation values and current period utilities for different realizations of the state.

To quantify the amount of distortion in the economy, Hansen and Sargent [2007] introduce measures of the conditional relative entropy between the reference and the distorted models. In particular, under the full information setting, the conditional relative entropy between the reference and the worst case model of the future state and signals evolution is defined as:

$$\epsilon_t^1(\phi_{t+1}, \xi_j) = \sum_{k=1}^{N} \int \tau(\xi_k, y_{t+1}|\xi_j)\phi(\xi_k, y_{t+1}) \log \phi(\xi_k, y_{t+1}) dy_{t+1}$$

and the conditional relative entropy between the reference and the worst case model of the current state by:

$$\epsilon_t^2(\psi_t) = \sum_{j=1}^{N} p_j \psi_j \log \psi_j.$$

The total conditional relative entropy between the reference model and the worst case model at date t is then given by:

$$\epsilon_t = \epsilon_t^2(\psi_t) + \tilde{\mathbb{E}}\left[\epsilon_t^1(\phi_{t+1}, z_t)\right|\mathcal{G}_t\left.\right] \equiv \epsilon_t^2(\psi_t) + \hat{\epsilon}_t^1(\phi_{t+1}). \tag{3.13}$$

In general, the recursion (3.9) does not have a closed-form solution. Instead, I look for a first order approximation to the representative agent's value function around the point $\theta^{-1} = 0$, which corresponds to the solution under the reference model. Notice that the approximation I construct here is different in its nature from the small noise approximations constructed in Campi and James [1996] and Anderson et al. [2010] as, instead of approximating the reference model value function around the deterministic steady state, I approximate around the value function corresponding to the zero signal precision case.[4] The following result holds:

Lemma 3.2. *The first order approximation to the value function around the point $\theta^{-1} = 0$ is given by:*

$$J(\pi; \theta^{-1}) = J_0(\pi) + \theta^{-1} J_1(\pi) + O_2(\theta^{-1}), \tag{3.14}$$

where $J_0(\pi)$ is the reference model value function and $J_1(\pi)$ is the first order derivative of the value function around $\theta^{-1} = 0$. In the non-default states of the economy, the first order approximation to J_0 and J_1 in terms of log-deviations, $\hat{\pi}$, from the stationary distribution $\overline{\pi}$ of the underlying Markov chain are given, respectively, by:

$$J_0(\pi) \approx \gamma_{00} + \gamma_{01}' diag(\overline{\pi}) \hat{\pi} \tag{3.15}$$

$$J_1(\pi) \approx \gamma_{10} + \gamma_{11}' diag(\overline{\pi}) \hat{\pi}, \tag{3.16}$$

where the coefficients γ_{00}, γ_{01}, γ_{10} and γ_{11} solve the system (C.2)-(C.6). The first order approximation to the implied distortion to the conditional joint distribution of next period's signals and state is then given by:

$$\phi_t(z^*, y^* | z = \xi_j) = 1 + \theta^{-1} \left(\varphi_{0j} + \varphi_{\pi,j}' diag(\overline{\pi}) \hat{\pi} + \varphi_{y,j}' \log f(y^*) \right), \tag{3.17}$$

where the constant coefficients φ_0, φ_π and φ_y satisfy (C.8)-(C.10), and the first order approxi-

[4]When the signal precision approaches 0, the agent is does not update the conditional probability distribution and, hence, the stationary distribution can be used as the prior distribution.

mation to the implied distortion to the conditional distribution of the current state by:

$$\psi_t(\xi_j) = 1 + \theta^{-1}\left[\zeta_{0j} + \zeta_{1j}' diag(\overline{\pi})\hat{\pi}\right], \tag{3.18}$$

where the coefficients ζ_0 and ζ_1 satisfy (C.11)-(C.12).

Proof. See Appendix C.1. □

Notice that J_0 is the value function of the representative agent under the reference model. The vector γ_{01} captures the first-order dependence of the reference model value function on the conditional distribution of the hidden state. From (C.3) we can see that the right hand side of the equation determining $\gamma_{01,j}$ is positive when the agent's utility in state j is higher than the stationary probability distribution weighted average of utility in different states. Thus, $\gamma_{01,j}$ is positive for states that have higher utility and negative for lower utility states. Intuitively, the expected present value of the representative agent's utility should be higher when the probability of the economy being in a good state is higher and lower when the probability of being in a bad state is higher. The value function derivative J_1 captures the first order depedence of the value function of the representative agent on the degree of her risk sensitivity. The fact that J_1 is time-varying implies that the agent's perceived risk attitudes change depending on the conditional distribution of the hidden state, with the vector γ_{11} describing the loadings on the individual components of the probability vector.

Consider now the implied distorted transition probablity matrix. Using (3.17), we have:

$$\begin{aligned} \tilde{\lambda}_{jk} &\equiv \tilde{\mathbb{P}}\left(z_{t+1} = \xi_k \mid z_t = \xi_j\right) \\ &\approx \lambda_{jk}\left[1 + \theta^{-1}\left(\varphi_{0j} + \varphi_{\pi,j}' diag(\overline{\pi})\hat{\pi} + \varphi_{y,j}'\Delta_k^1\right)\right]. \end{aligned} \tag{3.19}$$

Notice that this implies that, unlike the reference model, the transition probability matrix under the misspecified model is time-dependent. This effect introduces additional time variation into asset prices above that implied by the time evolution of fundamental asset values under the reference model.

3.3.1 Numerical example

To illustrate the impact of model misspecification on the representative investor's value function, I consider an example of a duopoly, where the (log) fundamental asset value evolves according to a five state Markov chain. In particular, assume that the two firms are symmetric in the coupon payment on their debt, C_i, the face value of their debt, D_i, the value of assets at default, a_{B_i}, cash-flows as a fraction of assets, δ_i, and the loadings on the aggregate component, ρ_i. I assume that the face value of the coupon bond is $D_i = 300$, serviced with the coupon rate $C_i = 6$. Each firm generates a fraction $\delta_i = 5\%$ of assets as cash flows and loading $\rho_i = 1$ of the log fundamental asset value on the aggregate asset value. At default, the value of assets of firm i is $a_{B_i} = C_i/\delta_i = 120$. The matrix ξ of possible values for the Markov state, z_t, is given by:

	Firm 1 (z_1)	Firm 2 (z_2)	Common (z_c)
State 1	$\log(2C/\delta)$	$\log(2C/\delta)$	$\log(2C/\delta$
State 2	$\log(C/\delta)$	$\log(2C/\delta)$	$\log(2C/\delta)$
State 3	$\log(C/\delta)$	$\log(C/\delta)$	$\log(C/\delta)$
State 4	0	$\log(C/\delta)$	$\log(C/\delta)$
State 5	$\log(C/\delta)$	0	$\log(C/\delta)$

Recall that the log fundamental asset value of firm i is given by $a_{it} = z_{it} + \rho_i z_{ct}$. Notice that state ξ_4 is the default state of firm 1 and the state ξ_5 is the default state of firm 2. The transition probability matrix, Λ, is chosen to have an equal probability of either staying in the same non-default state or transitioning to another non-default state and a much lower probability of transitioning to a default state:

$$\Lambda = \begin{bmatrix} 0.3125 & 0.3125 & 0.3125 & 0.0312 & 0.0312 \\ 0.3125 & 0.3125 & 0.3125 & 0.0312 & 0.0312 \\ 0.3125 & 0.3125 & 0.3125 & 0.0312 & 0.0312 \\ 0.0435 & 0.0435 & 0.0435 & 0.4348 & 0.4348 \\ 0.0435 & 0.0435 & 0.0435 & 0.4348 & 0.4348 \end{bmatrix}.$$

In Fig. 1 and Fig. 2, I plot the differences between the distorted transition probability matrix and the reference model transition probability matrix as a function of different combinations of the probabilities of state 1, p_1, and of state 2, p_2, for $\theta^{-1} = 0.5$ and $\theta^{-1} = 1$, respectively. Notice first that, as expected, under the misspecified model, the transition probabilities are shifted toward transitioning to default states. Further, as θ^{-1} increases, so that the agent becomes more misspecification-averse, this effect is larger. Next, notice that the effect is greater when the agent is more uncertain about the current state. That is, the greatest distortion to the transition probabilities occurs when (under the reference model) the agent has equal probabilities of being in any one of the non-default state. Intuitively, as the uncertainty about the state of the economy decreases under the reference model, the agent has relatively less misspecification concerns. Consider now the distorted probabilities over the current state, plotted in Fig. 3. Consistent with the intuition that the misspecification-averse agent tilts probabilities toward lower-utility states, we see that the probability of being in state 3 increase as misspecification aversion increases. Similarly to the case of transition probabilities, the impact of misspecification aversion is highest when the agent is most uncertain about which non-default state the economy is in.

3.4 Asset prices under the misspecified model

Since the representative agent evaluates expectations under the worst-case measure when making consumption decisions, the Euler equation holds under the worst-case measure. Therefore, assets can be priced using the Euler equation under the worst-case measure. In particular, under the misspecified model, the date t price of a claim to the equity of firm i satisfies:

$$V_{it} = \tilde{\mathbb{E}}\left[\delta_i A_{it} - C_i + \beta \mathbf{1}_{\tau_i > t+1} V_{i,t+1} \middle| \mathcal{G}_t\right]. \tag{3.20}$$

As with the value function, consider a first order expansion of the equity price around the reference model equity price. That is, I look for a first order approximation to the solution of the worst-case Euler equation (3.20) in the form:

$$V_i(\pi_t;\theta^{-1}) = V_{i0}(\pi_t) + \theta^{-1}V_{i1}(\pi_t) + O_2(\theta^{-2}). \tag{3.21}$$

The following result holds:

Lemma 3.3. *The first order approximations in terms of log-deviations from the stationary distribution to the value of equity of firm i under the reference model, V_{i0}, and to the derivative of the value equity of firm i, V_{i1}, are given, respectively, by:*

$$V_{i0}(\pi_t) = \nu_{i,00} + \nu'_{i,01}diag(\overline{\pi})\hat{\pi} + O_2(\hat{\pi}) \tag{3.22}$$

$$V_{i1}(\pi_t) = \nu_{i,10} + \nu'_{i,11}diag(\overline{\pi})\hat{\pi} + O_2(\hat{\pi}), \tag{3.23}$$

where the coefficients $\nu_{i,00}$, $\nu_{i,01}$, $\nu_{i,10}$ and $\nu_{i,11}$ solve the system of linear equations in Appendix C.2.

Proof. See Appendix C.2. □

Consider now the CDS spreads under the misspecified model. Recall that the spread on a default swap with maturity $T = t + 6n$ on the primary debt of firm i is given by:

$$c_i(t,T) = \frac{2X\tilde{\mathbb{E}}\left[\beta^{\tau_i - t}\mathbf{1}_{\tau_i<T}\middle|\mathcal{G}_t\right]}{\sum_{k=1}^{n}\beta^{6s}\tilde{\mathbb{E}}\left[\mathbf{1}_{\tau_i\geq t+6s}\middle|\mathcal{G}_t\right]}. \tag{3.24}$$

The misspecification concerns of the representative agent influence the CDS calculations in two ways. First, the conditional probability of the current state is distorted using $\psi(\xi_j)$, with the lower utility states receiving a higher probability. Second, the transition probability matrix Λ is replaced with the time-dependent distorted probability matrix $\tilde{\Lambda}$. Introduce the following notation:

$$\begin{aligned}\Upsilon(\pi,t,T) &= \tilde{\mathbb{E}}\left[\beta^{\tau_i - t}\mathbf{1}_{\tau_i<T}\middle|\mathcal{G}_t\right]\\ \Psi(\pi,t,T) &= \tilde{\mathbb{E}}\left[\mathbf{1}_{\tau_i\geq T}\middle|\mathcal{G}_t\right],\end{aligned}$$

so that:

$$c_i(t,T) = \frac{2X_i \Upsilon(\pi,t,T)}{\sum_{k=1}^{n} \beta^{6s} \Psi(\pi,t,t+6s)}.$$

By definition:

$$\begin{aligned}
\Upsilon(\pi,t,T) &= \sum_{j=1}^{N} p_{jt}\psi(\xi_j) \sum_{s=1}^{T-t} \beta^s \tilde{\mathbb{E}}\left[\mathbf{1}_{\tau_i = t+s} \middle| z_t = \xi_j\right] \\
&= \sum_{j=1}^{N} p_{jt}\psi(\xi_j) \sum_{s=1}^{T-t} \beta^s \tilde{\mathbb{P}}\left(\tau_i = t+s \middle| z_t = \xi_j\right) \\
\Psi(\pi,t,T) &= \sum_{j=1}^{N} p_{jt}\psi(\xi_j) \left(1 - \sum_{s=1}^{T-t} \tilde{\mathbb{E}}\left[\mathbf{1}_{\tau_i = t+s} \middle| z_t = \xi_j\right]\right) \\
&= \sum_{j=1}^{N} p_{jt}\psi(\xi_j) \left(1 - \sum_{s=1}^{T-t} \tilde{\mathbb{P}}\left(\tau_i = t+s \middle| z_t = \xi_j\right)\right).
\end{aligned}$$

Let $\tilde{q}_{ij}^n = \tilde{\mathbb{P}}\left(z_{t+n} = \xi_j,\ z_{t+n-1} \neq \xi_j, \ldots, z_{t+1} \neq \xi_j \middle| z_t = \xi_i\right)$ be the probability that $t+n$ is the first hitting time of state j conditional on being in state i at date t, so that $\tilde{\mathbb{P}}\left(\tau_i = t+n \middle| z_t = \xi_j\right) = \tilde{q}_{ji^*}^n$. Introduce also the matrix of the first hitting time probabilities: $\tilde{Q}^n = \left\{\tilde{q}_{ij}^n\right\}_{i,j=1}^{N}$ and let $\tilde{Q}_0^n = diag(\tilde{Q}^n)$. Then, from Seneta [1981], $\tilde{Q}^n$ is given recursively by:

$$\tilde{Q}^n = \tilde{\Lambda}\tilde{Q}_0^n.$$

Notice that, since $\tilde{\Lambda}$ is time-dependent, the first hitting time distribution under the misspecified model is also time-dependent.

Consider again the numerical example of Section 3.3.1. In Fig. 4, I plot the 5 year CDS rates for different values of θ^{-1} when no defaults have occurred. There are two effects of increasing misspecification aversion. First, as the agent becomes more misspecification averse, the CDS spreads increase since, under the misspecified model, the agent perceived a greater probability of transition to a default state. Second, notice that, for a given value of θ^{-1}, the CDS rate decreases as the agent becomes more certain about the state. As the agent becomes more misspecification averse, this effect increases. Intuitively, as the agent becomes more sure about the underlying state, she has less misspecification concerns about the conditional distribution of the current state, reducing the impact of misspecification concerns.

In Fig. 5, I consider also the percentage increase in the 5 year CDS rate of one firm after the other firm defaults. Since the more misspecification-averse representative agent places a higher probability on the third state, the percentage increase in CDS rates is smaller for higher values of θ^{-1}. That is, since the misspecification averse agent already has pessimistic views about the future, observing the default of one firm does not decrease her expectation about the time to default of the other firms as much as it does for the representative agent under the reference model.

Consider now the contagion effect of the default of firm i at date τ_i on the expected time to default of the surviving firms. I will use the following property of Markov chains:

Lemma 3.4. *Let $\{X_t\}_{t\geq 0}$ be a Markov chain on the probability space $(\Omega, \mathcal{F}, \mathbb{P})$ with transition matrix P and state space I. Define the hitting time of a subset A of I to be the random variable $\tau_A : \Omega \to \{0, 1, 2, \ldots\}$ such that $\tau_A(\omega) = \inf\{t \geq 0 : X_t(\omega) \in A\}$. Denote by $k_A(i)$ the mean time taken for $(X_t)_{t\geq 0}$ to reach A after starting from state i:*

$$k_A(i) = \mathbb{E}\left[\tau_A \middle| X_0 = i\right].$$

Let Q denote the matrix obtained by deleting the rows and columns corresponding to the set A from P. Then:

$$k_A(i) = \sum_{j\notin A}(I - Q)^{-1}_{ij} \tag{3.25}$$

Proof. See e.g. Seneta [1981]. □

Similarly to Frey and Schmidt [2010], I define the contagion effect as the change in the expected time to default of firm j at the default time τ_i of firm i, which is given by:

$$\hat{k}_{a_{B_j}}(\tau_i) - \hat{k}_{a_{B_j}}(\tau_i - 1) \equiv \tilde{\mathbb{E}}\left[k_{a_{B_j}} \middle| \mathcal{G}_{\tau_i}\right] - \tilde{\mathbb{E}}\left[k_{a_{B_j}} \middle| \mathcal{G}_{\tau_i - 1}\right]. \tag{3.26}$$

Notice that there are two opposing effects of observing one of the firms default. On the one hand, it reveals to the representative agent the current state of the economy, thus reducing the misspecification concerns faced by the agent. On the other hand, the conditional probability of the other firms defaulting next period increases as the agent knows that the aggregate component of the fundamental value is in its lowest state.

Consider once again the numerical example of Section 3.3.1. In Fig. 6, I plot the expected time to default for different values of the misspecification parameter, θ^{-1}. Notice that, as θ^{-1} increases, the expected time to default overall decreases. Intuitively, as the agent's misspecification aversion increases, the probability of transitioning to a default state increases, decreasing the expected time to default. In Fig. 7, I plot the percentage change in expected time to default after the other firm defaults. Since the more misspecification-averse representative agent places a higher probability on the third state, the percentage decrease in expected time to default decreases less for higher values of θ^{-1}. That is, since the misspecification averse agent already has pessimistic views about the future, observing the default of one firm does not decrease her expectation about the time to default of the other firms as much as it does for the representative agent under the reference model.

4 The 2007–2008 Financial Crisis

In this section, I calibrate the model in Section 3 using data on financial institutions. Although the returns on the equity of financial institutions accounts for a small portion of the overall level of consumption in the economy, these institutions were at the forefront of the 2007 financial crisis and, thus, to understand the asset price movements during the crisis, it is important to understand the movements in the prices of claims on these institutions. I begin by estimating the parameters of the reference model using the observations of book equity of financial institutions as firm-specific signals and the Case-Schiller 10 Cities Housing (CS10) Index as the signal about the common component of the asset values. I choose observations of the CS10 Index as the aggregate signal to capture the exposure of the financial institutions to risks associated with the national housing market. Notice from Table 1 that the financial institutions considered have higher correlations with the CS10 Index than with stock market indices, such as the S&P 500 index.

Next, I use historical observations of CDS spreads on financial institutions prior to the start of the crisis to estimate a value for θ – the model misspecification preference parameter. Using this estimate, I compute the implied relative entropy between the reference and the worst-case models and decompose the entropy calculation into the contributions from misspecification of the signal model and misspecification of the fundamental value of assets model for the whole time series. Finally, I compare the model-implied CDS rates and equity prices to the observed time series.

4.1 Estimating the reference model

To estimate the reference model, I use historical observations of the firm-specific signals and the aggregate signal. Below, I provide the outline of the estimation procedure. The details of the estimation are provided in Appendix A.

As observations of firm-specific signals, I use balance sheet data from COMPUSTAT. In particular, I use observations of book equity as the accounting signals. As observations of the aggregate signals, I take the time series of the Case-Schiller 10 Cities index. Notice that, while balance sheet data are observed at a quarterly frequency only, observations of the CS10 Index are available at a monthly frequency. The procedure described in Appendix A accounts explicitly for this dual frequency of observations. Notice that, to estimate the reference model, I only use observations up to Q2 2007 to avoid introducing measurement error by including observations of the accounting signal which reflect mark-downs taken since the start of the crisis, as well as the increased ambiguity discount in credit derivatives held on the balance sheets of these institutions.

Recall that the reference model is described by the parameters:

- Λ, $\{\xi_j\}_{j=1}^{N}$: transition probability matrix and states of the fundamental asset values

- $\{\rho_i\}_{i=1}^I$: firm-specific loadings on the common component
- Σ_u: signal error covariance matrix
- $\{D_i, C_i\}_{i=1}^I$: level of perpetual debt and coupon payments
- $\{\delta_i\}_{i=1}^I$: fraction of assets generated as cash-flows.

I begin the estimation by identifying the face value of the perpetual bond issued by firm i, D_i, with the last pre-crisis (Q2 2007) observation of the firm's value of long-term debt; notice that, since the model-implied debt has infinite maturity, long-term debt is a better measure than total debt as it excludes short-term liabilities. The coupon payment, C_i, is then chosen to make the level of debt D_i optimal. Following the model assumption that each firm generates a constant fraction δ_i of assets as cash-flows, I identify δ_i as the time-series average of the total earnings as a fraction of total assets. The rest of the parameters are identified using the Gibbs sampling procedure of Appendix A.

The estimated reference model parameters are presented in Tables 2 – 3 and the filtered time series of the expected fundamental value of each firm's assets under the reference model is plotted in Fig. 9. Notice first that the filtered fundamental asset value grows less over the considered time period than the observed book values and CS10 index. Intuitively, since the estimation extracts the filtered mean of the fundamental asset values from the observations of book values, the filtered values should have a smaller time trend. Notice also that the states of the economy are highly persistent. The probability of the firm-specific component of asset values staying in the same state next period is around 54% and the probability of the aggregate component of asset values staying in the same state next period is around 99%. Notice also that, although book values have a high correlation with the Case Schiller 10 Index, the estimated firm-specific loadings on the aggregate component of asset values is lower, ranging from 27% for JP Morgan and less than 1% for Bear Stearns. Finally, notice that the signal errors have low cross-correlations of at most 7% and a higher variance, ranging from 73% for the Case Schiller 10 Index to 15% for Goldman Sachs.

4.2 Estimating the misspecification preference parameter

Consider now estimating the misspecification preference parameter, θ^{-1}. Rewrite the CDS equation (3.24) as:

$$0 = 2X_i \Upsilon(\pi, t, T) - c_i(t, T) \sum_{k=1}^{n} \beta^{6s} \Psi(\pi, t, t + 6s).$$

To estimate θ^{-1}, I assume that the CDS rates are observed with a measurement error. In particular, assume that the observation equation is given by:

$$0 = 2X_i \Upsilon(\pi, t, T) - \hat{c}_i(t, T) \sum_{k=1}^{n} \beta^{6s} \Psi(\pi, t, t + 6s) + \epsilon_{iT,t},$$

where $\hat{c}_i(t, T)$ are the observed CDS rates and the vector of maturity-specific measurement errors $\epsilon_{it} = [\epsilon_{i1,t}, \ldots, \epsilon_{iT,t}]$ is normally distributed and i.i.d. across time and firms: $\epsilon_i \sim N(0, \Sigma_\epsilon)$. I make draws of θ^{-1} using a Random Walk Metropolis algorithm with a flat prior. The accept/reject probability for the draws of θ^{-1} is the ratio of the likelihood of the CDS rates for all firms, at all available data points and for all available maturities.

I conduct three estimations of the parameter θ^{-1} using different data sub-periods: the pre-crisis period, the period from the start of the crisis to the bailout of Bear Stearns and the period from the bailout of Bear Stearns to the liquidation of Lehman Brothers. The results of these estimations are presented in Table 4. Notice that the three different periods do not yield significantly different estimates of θ^{-1}, suggesting that the investors attitudes toward model misspecification do not change significantly during the crisis period. Instead, the observed changes in CDS rates are due to shifts in what kind of misspecification the agents are more concerned about.

Consider now the model-implied time series evolution of credit spreads. Since the estimate of θ^{-1} does not change during the crisis, I use the pre-crisis estimate of θ^{-1} to compute the model-implied CDS spreads. Table 5 presents the observed 5 year CDS spreads for the financial institutions at five dates of interest – before the start of the crisis, July 2007, at the start of the crisis in August 2007, after the bailout of Bear Stearns in March 2008, after the liquidation of Lehman Brothers in September 2008 and after the introduction of TARP in October 2008 – together with the model-implied CDS rates at these dates. Notice first that the model-implied CDS rates follow the observed pattern of increasing during the financial crisis. Further, for most institutions, the implied CDS rates match the levels of CDS spreads over time, although the performance of the estimated model worsens after the liquidation of Lehman Brothers.

To evaluate how misspecification shifts during the crisis, I compute the entropy contributions from misspecification of the joint distribution of next period's signals and state and misspecification of the conditional probability distribution of the current state. Notice first that the entropy contribution from the misspecification of the conditional probability distribution of the current state is much smaller than the contribution from the misspecification of the joint distribution of next period's signals and state. Notice also that both components of entropy increase after the BNP Paribas announcement in August 2007, which is consistent with the Caballero and Krishnamurthy [2008a] intuition that ambiguity increased during the crisis. The period between the bailout of Bear Stearns and the liquidation of Lehman Brothers in September 2008 on the other hand only lead to increase in the entropy from the misspecification of the conditional probability distribution of the current state. Intuitively, while investors observed the bailout of Bear Stearns, they were not sure that the government would conduct any more bailouts, increasing the overall uncertainty in the economy.

To further understand the time series evolution of asset prices during the crisis, consider the time series evolution of expected time to default of the financial institutions. Table 6 presents the expected time to default for the financial institutions at five dates of interest – before the start of the crisis, July 2007, at the start of the crisis in August 2007, after the bailout of Bear Stearns in March 2008, after the liquidation of Lehman Brothers in September 2008 and after the introduction of TARP in October 2008 – together with the percentage change in the time to default relative to the previous month. The initial BNP Paribas announcement in August 2007 only lead to a decrease in the expected time to default for only Bank of America and Goldman

Sachs. The bailout of Bear Stearns lead to a decrease in expected time to default of all firms, except for Bank of America. Notice, however, that while the increase in expected time to default for Bank of America was small – only 2% , the decrease in expected time to default for the other firms was much larger – from 16% for Morgan Stanley to 32% for JP Morgan. Thus, overall, the bailout of Bear Stearns induced contagion effects on the rest of the financial institutions, with the degree of contagion varying across firms. Similarly, the liquidation of Lehman Brothers lead to a decrease in the expected time of default for all firms except Bank of America and Goldman Sachs. It is important to note that, for these two institutions, the increase in expected time to default was significant – 20% for Bank of America and 120% for Goldman Sachs.

Compare this to the evolution of expected times to default under the reference model, presented in Panel B of Table 6. Notice first that, under the reference model, the expected time to default is longer at all dates than under the misspecified model. Intuitively, the misspecification-averse agent perceives the probability of default next period to be greater than under the reference model, decreasing the expected time to default. Next, consider the time evolution of the expected times to default. The initial BNP Paribas announcement leads to a slight (0.03%) decrease in the expected time to default for each of the institutions. As under the misspecified model, the decrease in the expected time to default is much greater after the bailout of Bear Stearns. Notice, however, that the decrease under the reference model is greater than under the misspecified model. Intuitively, since the misspecification-averse agent already has more pessimistic views of the future, observing the bailout of Bear Stearns did not have as a large of an impact on her beliefs as it did on the beliefs under the reference model. Similarly, the decrease under the reference model after the liquidations of Lehman Brothers is larger than under the misspecified model. After introduction of TARP in October 2008, however, the increase in the expected time to default under the reference model is much smaller than under the misspecified model.

Finally, consider the implied time series evolution of equity prices during the crisis. Fig. 11 plots the observed evolution of equity prices together with the model-implied evolution.

Although we cannot hope to match the level of equity prices since the firm earnings model in the paper is extremely simplistic, the model should be able to match the observed movements in equity prices. Comparing the model-implied evolution to the true evolution of equity prices, we see that the model-implied equity prices lag the observed equity prices. This is not surprising since the signals used to construct the time series evolution of conditional probabilities are backward-looking; for example, the Case-Schiller 10 Index is constructed using observations over the previous three months. The model is able to capture the overall downward trend of equity prices during the crisis and especially well the sharp drop in equity prices after the bailout of Bear Stearns and after the liquidation of Lehman Brothers.

5 Conclusion

In this paper, I consider the implications of model misspecification for default swap spreads. Using an incomplete information version of the Black and Cox [1976] model of credit spreads as the reference model, I find that introducing misspecification concerns exacerbates the imperfect information problem faced by the representative agent. This leads to an increased level of default swap spreads overall and greater sensitivity of CDS spreads to bad news. The misspecification-averse agent perceives the probability of default next period to be higher than under the reference model, increasing CDS spreads and decreasing expected time to default. Observing a bad signal not only increases the conditional probability of being in a low-payoff state in the current period but also increased the perceived probability of default in the next period.

To investigate whether the model can produce reasonable magnitudes of CDS spreads, I estimated the parameters of the reference model using observations of the book value of equity of several financial institutions as firm-specific signals and of Case Schiller 10 Index as observations of aggregate signals. The misspecification preference parameter θ^{-1} was then estimated using observations of CDS spreads for the financial institutions over time. The results of the estimation procedure suggest that, while agents' preference toward model misspecification did not change during the crisis, the amount of entropy in the economy and how that entropy is decomposed into the contributions from misspecification of the joint distribution of next period's signals and state and misspecification of the conditional probability distribution of the current state did change. In particular, the initial BNP Paribas announcement in August 2007 lead to an increase in both components of entropy and, especially, in the entropy from from misspecification of the joint distribution of next period's signals and state. The period between the bailout of Bear Stearns and the liquidation of Lehman Brothers in September 2008 on the other hand only lead to increase in the entropy from the misspecification of the conditional probability distribution of the current state.

Examining the implied time-series evolution of equity prices, I find that, while the model is able to match the overall movements of the equity prices, the model-implied equity prices lag the observed equity prices. A possible avenue of future research is to estimate the model under the assumption that the representative agent in the economy observes more information than the econometrician estimating the model. While allowing for equity prices to adjust quicker to news in the market, this would also allow us to estimate the model at a higher frequency and, thus, extract more information from the observed CDS spreads.

Notice that, while the model described in this paper is geared toward explaining the observed increases in CDS spreads, similar intuition could be used to explain observed changes to prices of collateralized debt obligations (CDOs) and other complex securities. In fact, since arguably CDOs have a more complicated underlying structure than default swaps, model misspecification concerns would be even more relevant in pricing these securities. A formal treatment of this problem, however, is left for future research.

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A Reference Model Estimation

I estimate the time series parameters of the reference model using a Gibbs sampling procedure. Recall the Gibbs sampling allows to sequentially make parameter draws from conditional posteriors. Because the model can be broken down into as many conditional posteriors as needed, it is possible to fully estimate the reference model specified in Section 3.

Recall that for the model of Section 3.1, the observations in the economy are given by:

$$\begin{aligned} y_{it} &= \xi_{i,s_t} + \rho_i \xi_{I+1,s_t} + u_{it}, \qquad i = 1, \ldots, I \\ y_{ct} &= \xi_{I+1,s_t} + u_{ct}, \end{aligned}$$

where s_t is the indicator of the state at date t. In reality, observations of book value occur at a quarterly frequency while observations of the aggregate signal occur at a monthly frequency. Denote by n_y the number of periods in between observations of firm-specific signals and by t_y the number of available observations of the firm-specific signals.

To reduce the number of parameters to be estimated, I impose additional restrictions on the model. In particular, I assume that the vector of the firm-specific components $z_{ft} \equiv [z_{1t}, \ldots, z_{It}]$ of the hidden state vector z evolves independently of the aggregate component, z_{ct}. That is, I assume that z_{ft} and z_{ct} evolve as two independent Markov chains. The vector of firm-specific components z_f evolves as an n_f-state Markov chain, with values $\zeta_{f1}, \ldots, \zeta_{f,n_f}$ and transition probability matrix Ω_f defined by:

$$\{\Omega_f\}_{jk} \equiv \omega_{f,ik} = \mathbb{P}\left(z_{f,t+1} = \zeta_{fk} \middle| \, z_{ft} = \zeta_{fj} \right).$$

Similarly, the aggregate component z_{ct} evolves as an n_c Markov chain, with values $\zeta_{c1}, \ldots, \zeta_{cn_c}$ and transition matrix Ω_c defined by:

$$\{\Omega_c\}_{jk} \equiv \omega_{ci,jk} = \mathbb{P}\left(z_{c,t+1} = \zeta_{ck} \middle| \, z_{ct} = \zeta_{cj} \right).$$

I impose also the assumption that the signal errors of the aggregate signal are uncorrelated with the signal errors of the firm-specific signals but allow for the errors of the firm-specific signals to be cross-sectionally correlated. That is, I partition the signal covariance matrix into:

$$\Sigma_u = \begin{bmatrix} \Sigma_{uf} & \vec{0}_{I,1} \\ \vec{0}_{1,I} & \Sigma_{uc} \end{bmatrix},$$

where Σ_{uf} is the covariance matrix of the firm-specific signals and Σ_{uc} is the variance of the aggregate signal. Notice that this formulation allows me to estimate the firm value at default directly from the signal observations: since z_{ft} and z_{ct} evolve as two independent Markov chains, it is possible to recover the lowest value of the firm-specific component and the lowest value of the common component without observing default.

Denote by Θ the full set of parameters to be estimated:

$$\Theta = \left\{ \Omega_f, \Omega_c, \zeta_{f1}, \ldots, \zeta_{fn_f}, \zeta_{c1}, \ldots, \zeta_{cn_c}, \rho_1, \ldots, \rho_I, \Sigma_{uf}, \Sigma_{uc}, \{s_{ft}\}_{t=1}^{t_y}, \{s_{ct}\}_{t=1}^{T} \right\}$$

and by Θ_{-A} the set of all parameters except A: $\Theta_{-A} = \Theta \setminus A$. With the above assumptions, the main steps in the Gibbs procedure are as follows:

Step 1. Conditional on a draw of $\Theta_{-\Omega_f}$, make a draw of Ω_f

Step 2. Conditional on a draw of $\Theta_{-\Omega_c}$, make a draw of Ω_c

Step 3. Conditional on a draw of $\Theta_{-\{\zeta_{ci}\}_{i=1}^{n_c}}$, make a draw of $\zeta_{c1}, \ldots, \zeta_{c,n_c}$

Step 4. Conditional on a draw of $\Theta_{-\left\{\zeta_{fi}\right\}_{i=1}^{n_f}}$, make a draw of $\zeta_{f1}, \ldots, \zeta_{f,n_f}$

Step 5. Conditional on a draw of $\Theta_{-\{\rho_i\}_{i=1}^{I}}$, make a draw of $\rho_1, \ldots, \rho_I$

Step 6. Conditional on a draw of $\Theta_{-\Sigma_{uf}}$, make a draw of Σ_{uf}

Step 7. Conditional on a draw of $\Theta_{-\Sigma_{uc}}$, make a draw of Σ_{uc}

Step 8. Conditional on a draw of $\Theta_{-\{s_{ct}\}_{t=1}^{T}}$, make a draw of $\{s_{ct}\}_{t=1}^{T}$

Step 9. Conditional on a draw of $\Theta_{-\left\{s_{ft}\right\}_{t=1}^{t_y}}$, make a draw of $\{s_{ft}\}_{t=1}^{t_y}$

Step 10. Permute the state indicators

For the conditional posteriors below, I rely on Gibbs sampling results for regime-switching models. The initial application of MCMC estimation methods to regime-switching models in due to Albert and Chib [1993], who estimate an autoregressive model with Markov jumps following a two-state Markov process. McCulloch and Tsay [1994] extend this to situations where the regime-switching model includes non-regime-specific (common) components. For the most part (and in the algorithm below), practical MCMC estimation uses the principle of data augmentation and treats the indicators of the state of the latent Markov chain as missing data. Treating observations of the latent Markov chain as missing data allows for the use of conjugate priors in estimating the parameters of the model. For a more exhaustive discussion of the use of MCMC methods for estimating the parameters of Markov chains, see Fruhwirth-Schnatter [2006].

A.1 Conditional on a draw of $\Theta_{-\Omega_f}$, make a draw of Ω_f

Denote by $\tilde{\Omega}_f$ the n_y-periods-ahead transition probability matrix of the firm-specific value vector z_{ft}: $\tilde{\Omega}_f = \Omega_f^{n_y}$. Since obsevations of the firm-specific signals occur only every n_y periods and the aggregate signals are not informative about the firm-specific state, I make draws of $\tilde{\Omega}_f$ and then infer the corresponding draw of Ω_f. Generalizing the results of Albert and Chib [1993] and McCulloch and Tsay [1994] to the multiple state case, the conjugate prior for the j^{th} row of $\tilde{\Omega}_f$ is :

$$\tilde{\omega}_{f,j} \sim Dir\left(\alpha_{j1}^{f}, \ldots, \alpha_{j,n_f}^{f}\right),$$

where Dir denotes the Dirichlet distribution[5]. The posterior is then given by:

$$\tilde{\omega}_{f,j} \sim Dir\left(\alpha_{j1}^{f} + n_{j1}^{f}, \ldots, \alpha_{j,n_f}^{f} + n_{jn_f}^{f}\right),$$

[5] Recall that the Dirichlet distribution generalizes the beta distribution to the multinomial case

where n_{jk}^{f} is the number of times the Markov chain z_f transitions from state j to state k in the current draw of $\{s_{ft}\}_{t=1}^{t_y}$. Once a draw of $\tilde{\Omega}_f$ is made, the corresponding draw of the original transition probability matrix is computed as $\Omega_f = \tilde{\Omega}_f^{\frac{1}{n_y}}$.

A.2 Conditional on a draw of $\Theta_{-\Omega_c}$, make a draw of Ω_c

Similarly to $\tilde{\Omega}_f$, the conjugate prior for the j^{th} row of Ω_c is the Dirichlet distribution:

$$\omega_{c,j} \sim Dir\left(\alpha_{j1}^{c}, \ldots, \alpha_{j,n_c}^{c}\right),$$

and the posterior is given by:

$$\omega_{c,j} \sim Dir\left(\alpha_{j1}^{c} + n_{j1}^{c}, \ldots, \alpha_{j,n_c}^{c} + n_{j,n_c}^{c}\right),$$

where n_{jk}^{c} is the number of times the Markov chain z_c transitions from state j to state k in the current draw of $\{s_{ct}\}_{t=1}^{T}$.

A.3 Conditional on a draw of $\Theta_{-\{\zeta_{ci}\}_{i=1}^{n_c}}$, make a draw of $\zeta_{c1}, \ldots, \zeta_{c,n_c}$

To derive the conditional posterior of ζ_{ci}, $i = 1, \ldots, n_c$, notice that the firm-specific signals contain information about the common component of the fundamental asset values. In particular, notice that the likelihood function of the signals is given by:

$$\begin{aligned}
\mathcal{L}\left(y \middle| \Theta\right) \;\propto\; & \exp\left\{-\frac{1}{2}\sum_{\tau=1}^{t_y}\left(y_{f,\tau\Delta_y} - \zeta_{f,s_{f,\tau\Delta_y}} - \rho\zeta_{c,s_{c,\tau\Delta_y}} - \overline{u}_f\right)' \Sigma_{uf}^{-1}\left(y_{f,\tau\Delta_y} - \zeta_{f,s_{f,\tau\Delta_y}} - \rho\zeta_{c,s_{c,\tau\Delta_y}} - \overline{u}_f\right)\right\} \\
\times\; & \exp\left\{-\frac{1}{2}\sum_{t=1}^{T}\frac{\left(y_{ct} - \zeta_{c,s_{ct}} - \overline{u}_c\right)^2}{\Sigma_{uc}}\right\}
\end{aligned}$$

Let $i_1 < i_2 < \ldots < i_{n_j}$ denote all the time indices such that $s_{ci_k} = j$ and let $y_{c,j} = \left(y_{ci_1}, \ldots, y_{ci_{n_j}}\right)'$. The conjugate prior is given by:

$$\zeta_{cj} \sim N\left(\zeta_{c,0j},\ \Sigma_{uc} A_{c,0j}^{-1}\right)$$

and the conditional prior by:

$$\zeta_{cj} \sim N\left(\bar{\zeta}_{cj},\ \Sigma_{uc}\bar{A}_{cj}^{-1}\right),$$

where

$$\begin{aligned}
\bar{A}_{cj} &= n_{y_j}\Sigma_{uc}\rho'\Sigma_{uf}^{-1}\rho + n_j + A_{c,0j} \\
\bar{\zeta}_{cj} &= \bar{A}_{cj}^{-1}\left[\Sigma_{uc}\rho'\Sigma_{uf}^{-1}\sum_{k=1}^{n_{y_j}}\left(y_{f,\tau_k\Delta_y} - \zeta_{f,s_{1\tau_k\Delta_y}} - \bar{u}_f\right) + \sum_{k=1}^{n_j}\left(y_{c,i_k} - \overline{u}_c\right) + A_{c,0j}\zeta_{c,0j}\right]
\end{aligned}$$

A.4 Conditional on a draw of $\Theta_{-\left\{\zeta_{fi}\right\}_{i=1}^{n_f}}$, make a draw of $\zeta_{f1}, \ldots, \zeta_{f,n_f}$

Denote: $\tilde{y}_{ft} = y_{ft} - \rho\zeta_{cs_{ct}} - \bar{u}$. Let $j_1 < j_2 < \ldots < j_{n_j}$ denote all the time indices such that $s_{fj_k} = j$ and let $\tilde{y}_{fj} = \left(\tilde{y}_{f,j_1}, \ldots, \tilde{y}_{f,j_{n_j}}\right]'$. Then the conjugate prior distribution is:

$$\zeta_{fj} \sim N\left(\zeta_{f,0j},\ \Sigma_{uf} A_{f,0j}^{-1}\right)$$

and the conditional posterior by:

$$\zeta_{fj} \sim N\left(\bar{\zeta}_{fj},\ \Sigma_{uf} \bar{A}_{fj}^{-1}\right),$$

where:

$$\begin{aligned} \bar{A}_{fj} &= n_j + A_{f,j0} \\ \bar{\zeta}_{fj} &= \bar{A}_{fj}^{-1}\left[\sum_{k=1}^{n_j} \tilde{y}_{f,j_k} + A_{f,j0}\zeta_{f,j0}\right] \end{aligned}$$

A.5 Conditional on a draw of $\Theta_{-\{\rho_i\}_{i=1}^{I},\Sigma_{uf}}$, make a draw of $\rho_1, \ldots, \rho_I$ and

Define: $\tilde{y}_{ft} = y_{ft} - \overline{u}_f - \zeta_{fs_{ft}}$. The prior distribution for the vector ρ is then

$$\rho \sim N\left(\beta_0, \Sigma_{uf} A_{0\rho}^{-1}\right)$$

and the conjugate posterior is given by:

$$\rho \sim N\left(\beta_*, \Sigma_{uf} A_{*,\rho}^{-1}\right),$$

where:

$$\begin{aligned} A_{*,\rho} &= \sum_{t=1}^{t_y} \zeta_{c,s_{ct}}^2 + A_{0\rho} \\ \beta_* &= A_{*,\rho}^{-1}\left(\sum_{t=1}^{t_y} \tilde{y}_{ft}\zeta_{c,s_{ct}} + A_{0\rho}\beta_0\right). \end{aligned}$$

The conjugate prior for Σ_{uf} is given by:

$$\Sigma_{uf} \sim IW(\nu, V)$$

and the posterior is given by:

$$\Sigma_{uf} \sim IW(\nu + t_y, V + S),$$

where

$$S = \sum_{t=1}^{t_y} \left(\tilde{y}_{f,t\Delta_y} - \rho \zeta_{c,s_{c,t\Delta_y}} \right)' \left(\tilde{y}_{f,t\Delta_y} - \rho \zeta_{c,s_{c,t\Delta_y}} \right) + (\beta_* - \beta_0)' A_{0\rho} (\beta_* - \beta_0).$$

A.6 Conditional on a draw of $\Theta_{-\Sigma_{uc}}$, make a draw of Σ_{uc}

In this case, the conjugate prior is an inverse χ^2-distribution:

$$\Sigma_{uc} \sim \frac{\nu_c \bar{s}_c^2}{\chi^2(\nu_c)}$$

and the conditional posterior by:

$$\Sigma_{uc} \sim \frac{\nu_c \bar{s}_c^2 + TE_c}{\chi^2(\nu_c + T)},$$

where:

$$TE_c = \sum_{t=1}^{T} (y_{ct} - \zeta_{cs_{ct}} - \overline{u}_c)^2.$$

A.7 Conditional on a draw of $\Theta_{-\{s_{ct}\}_{t=1}^{T}}$, make a draw of $\{s_{ct}\}_{t=1}^{T}$

Denote by S_{ct} the history of observations of the aggregate regime up to date t, $S_{c,-t}$ the full history of the aggregate regime except at date t. The conditional posterior is then given by:

$$\begin{aligned}
\mathbb{P}\left(s_{ct} | \{y_{c\tau}\}_{\tau=1}^{T}, S_{c,-t}\right) &\propto \mathbb{P}(s_{ct}|s_{c,t-1})\mathbb{P}(s_{c,t+1}|s_{ct}) \exp\left\{ -\frac{1}{2} \frac{(y_{ct} - \zeta_{c,s_{ct}})^2}{\Sigma_{uc}} \right\} \qquad t \neq n n_y, n \in \mathbb{N} \\
\mathbb{P}\left(s_{ct} | \{y_{c\tau}\}_{\tau=1}^{T}, S_{c,-t}\right) &\propto \mathbb{P}(s_{ct}|s_{c,t-1})\mathbb{P}(s_{c,t+1}|s_{ct}) \exp\left\{ -\frac{1}{2} \frac{(y_{ct} - \zeta_{c,s_{ct}})^2}{\Sigma_{uc}} \right\} \\
&\times \exp\left\{ -\frac{1}{2} \left(y_{ft} - \zeta_{f,s_{ft}} - \rho\zeta_{cs_{ct}}\right)' \Sigma_{uf}^{-1} \left(y_{ft} - \zeta_{f,s_{ft}} - \rho\zeta_{cs_{ct}}\right) \right\} \qquad t = n n_y, n \in \mathbb{N}
\end{aligned}$$

A.8 Conditional on a draw of $\Theta_{-\left\{s_{ft}\right\}_{t=1}^{t_y}}$, make a draw of $\{s_{ft}\}_{t=1}^{t_y}$

Denote S_{ft} the history of observations of the firm-specific regime up to date t, $S_{f,-t}$ the full history of observations of the firm-specific regime except at date t. Then, the conditional posterior for s_{ft} is given by:

$$\mathbb{P}\left(s_{ft} | \{y_{f\tau}\}_{\tau=1}^{T}, S_{f,-t}\right) \propto \mathbb{P}(s_{ft}|s_{f,t-1})\mathbb{P}(s_{f,t+1}|s_{ft}) \exp\left\{ -\frac{1}{2} \left(y_{ft} - \zeta_{f,s_{ft}} - \rho\zeta_{cs_{ct}}\right)' \Sigma_{uf}^{-1} \left(y_{ft} - \zeta_{f,s_{ft}} - \rho\zeta_{cs_{ct}}\right) \right\}.$$

A.9 Permute the state indicators

As discussed in Fruhwirth-Schnatter [2001], the behavior of the sampler described above is somewhat unpredictable, and the sampler might be trapped at one modal region of the Markov mixture posterior distribution or may jump occasionally between different model regions causing label switching. A simple but efficient solution to obtain a sampler that explores the full Markov mixture posterior distribution is suggested in Fruhwirth-Schnatter [2006]. Each draw from the Gibbs sampler is concluded by selecting randomly one of $n_f!$ possible permutations of the current labeling of the firm-specific states and one $n_c!$ possible permutations of the current labeling of the aggregate states. This permutation is then applied to the transition probability matrices Ω_f and Ω_c, the state-specific parameters ζ_f and ζ_c and the state indicators $s_f^{t_y}$ and s_c^T.

B Model Misspecification

In this section, I describe the derivation of the risk-sensitive recursion (3.9) and the distortions (3.11)-(3.12) to the filtering distributions. I rely on the results of Hansen and Sargent [2005], Hansen and Sargent [2007] to formulate the model misspecification problem faced by the representative investor.

Let M_t be a non-negative $\mathcal{F}_t$-measurable random variable, with $\mathbb{E}[M_t] = 1$. Using M_t as a Radon-Nikodym derivative generates a distorted probability measure that is absolutely continuous with respect to the probability measure over $\mathcal{F}_t$ generated by the model (3.1). Under the distorted measure, the expectation of a bounded $\mathcal{F}_t$-measurable random variable W_t is $\tilde{\mathbb{E}}[W_t] = \mathbb{E}[M_t W_t]$.

To construct the implied (distorted) conditional density, Hansen and Sargent [2007] factor the martingale M_t into one-step-ahead random variables. More specifically, for a non-negative martingale $\{M_t\}_{t\geq 0}$ form:

$$m_{t+1} = \begin{cases} \frac{M_{t+1}}{M_t} & \text{if } M_t > 0 \\ 1 & \text{if } M_t = 0 \end{cases}$$

Then $M_{t+1} = m_{t+1} M_t$ and, for any $t \geq 0$, the martingale M_t can be represented as:

$$M_t = M_0 \prod_{j=1}^{t} m_j,$$

where the random variable M_0 has unconditional expectation equal to unity. Notice that, by construction, m_{t+1} has date t conditional expectation equal to unity. Thus, for a bounded $\mathcal{F}_{t+1}$-measurable random variable W_{t+1}, the distorted conditional expectation implied by the martingale $\{M_t\}_{t\geq 0}$ is constructed as:

$$\tilde{\mathbb{E}}[W_{t+1}|\,\mathcal{F}_t] \equiv \frac{\mathbb{E}[M_{t+1}W_{t+1}|\,\mathcal{F}_t]}{\mathbb{E}[M_{t+1}|\,\mathcal{F}_t]} = \frac{\mathbb{E}[M_{t+1}W_{t+1}|\,\mathcal{F}_t]}{M_t} = \mathbb{E}[m_{t+1}W_{t+1}|\,\mathcal{F}_t],$$

provided that $M_t > 0$. Let $\mathcal{M}_t$ be the space of all non-negative $\mathcal{F}_t$-measurable random variables m_t for which $\mathbb{E}[m_t|\,\mathcal{F}_{t-1}] = 1$. The elements of $\mathcal{M}_{t+1}$ represent all possible distortions of the conditional distribution over $\mathcal{F}_{t+1}$ given $\mathcal{F}_t$; that is, each $m_{t+1} \in \mathcal{M}_{t+1}$ represents a possible distortion to the underlying asset value dynamics. The amount of distortion introduced by m_{t+1} is measured each period using the conditional relative entropy between the reference and distorted models:

$$\epsilon_t^1(m_{t+1}) = \mathbb{E}\left[m_{t+1} \log m_{t+1} \middle| \mathcal{F}_t \right]. \tag{B.1}$$

To introduce distortion to the signal model, consider factoring the martingale M_t in a different way. More specifically, introduce the $\mathcal{G}_t$-measurable random variable $\hat{M}_t = \mathbb{E}\left[M_t \middle| \mathcal{G}_t \right]$ and define:

$$h_t = \begin{cases} \frac{M_t}{\hat{M}_t} & \text{if } \hat{M}_t > 0 \\ 1 & \text{if } \hat{M}_t = 0 \end{cases}$$

The $\mathcal{G}_t$-measurable random variable $\hat{M}_t$ implies a likelihood ratio for the partial information set $\mathcal{G}_t$ while the $\mathcal{F}_t$-measurable random variable h_t represents distortions to the probability distribution over $\mathcal{F}_t$ given $\mathcal{G}_t$. Define $\mathcal{H}_t$ to be the space of all non-negative $\mathcal{F}_t$-measurable random variables h_t for which $\mathbb{E}\left[h_t \middle| \mathcal{G}_t \right] = 1$. Similarly to (B.1), the amount of distortion induced by h_t is measured as:

$$\epsilon_t^2(h_t) = \mathbb{E}\left[h_t \log h_t \middle| \mathcal{G}_t \right]. \tag{B.2}$$

To solve for the worst-case likelihood, introduce an entropy penalization parameter $\theta > 0$ which captures the beliefs of the representative agent about the amount of misspecification in the economy: as θ increases, the set of admissible alternative models decreases, with the limiting case $\theta = \infty$ corresponding to only the reference model being admissible. Begin by considering the full-information case. Corresponding to each $\mathcal{F}_{t+1}$-measurable random variable m_{t+1} is a relative density $\phi(z^*, y^*)$. The minimizing agent solves:

$$\min_{\phi \geq 0} \sum_{j=1}^{N} \int \left[W(\xi_j, \pi^*, y^*) + \theta \log \phi(\xi_j, y^*) \right] \phi(\xi_j, y^*) \tau(\xi_j, y^* | z, y) dy^*$$

subject to:

$$\sum_{j=1}^{N} \int \phi(\xi_j, y^*) \tau(\xi_j, y^* | z, y) dy^* = 1,$$

where $*$ denote next period values and $\tau(z^*, y^* | z, y)$ is the joint transition probability:

$$\tau(z^*, y^* | z, y) = |2\pi\Sigma_u|^{-\frac{1}{2}} \lambda_{zz^*} \exp\left[-\frac{1}{2}(y^* - z^* - \overline{u})' \Sigma_u^{-1} (y^* - z^* - \overline{u}) \right]. \tag{B.3}$$

The solution to the minimization problem implies a transformation T^1 that maps the value function that depends on next period's state (ξ_j, π^*, y^*) into a risk-adjusted value function that depends on the current state (z, π, y):

$$T^1(W | \theta) = -\theta \log \sum_{j=1}^{N} \int \exp\left(-\frac{W(\xi_j, \pi^*, y^*)}{\theta} \right) \tau(\xi_j, y^* | z, y) dy^*. \tag{B.4}$$

The minimizing choice of ϕ is given by:

$$\phi_t(z^*, y^*) = \frac{\exp\left(-\frac{W(z^*,\pi^*,y^*)}{\theta}\right)}{\mathbb{E}\left[\exp\left(-\frac{W(z^*,\pi^*,y^*)}{\theta}\right)\Big| \mathcal{F}_t\right]}.$$

Similarly, corresponding to each $\mathcal{G}_t$-measurable random variable h_t is a relative density $\psi(z)$, with the worst-case distortion given as the solution to:

$$\min_{\psi \geq 0} \sum_{j=1}^{N} \left[\hat{W}(\pi, \xi_j) + \theta \log \psi(\xi_j)\right] \psi(\xi_j) p_j$$

subject to:

$$\sum_{j=1}^{N} \psi(\xi_j) p_j = 1.$$

This implies another operator:

$$T^2(\hat{W}|\theta)(\pi) = -\theta \log \sum_{j=1}^{N} \exp\left(-\frac{\hat{W}(\pi, \xi_j)}{\theta}\right) \psi(\xi_j) p_j. \tag{B.5}$$

The corresponding minimizing choice of ψ is given by:

$$\psi_t(z) = \frac{\exp\left(-\frac{\hat{W}(\pi,z)}{\theta}\right)}{\mathbb{E}\left[\exp\left(-\frac{\hat{W}(\pi,z)}{\theta}\right)\Big| \mathcal{G}_t\right]}.$$

C Proofs

C.1 Proof of Lemma 3.2

To obtain a first order approximation to the value function around the point $\theta^{-1} = 0$, we need to take a second order expansion of the risk-sensitive recursion. In particular, approximate:

$$\begin{aligned}
\exp\left[-\frac{J(p;\theta^{-1})}{\theta}\right] &\approx 1 - J(p,0)\theta^{-1} + \frac{1}{2}\left[J(p;0)^2 - 2\frac{\partial J(p;0)}{\partial \theta^{-1}}\right]\theta^{-2} \\
\exp\left[-\frac{U(\xi_j) + \beta J(p;\theta^{-1})}{\theta}\right] &\approx 1 - \left[U(\xi_j) + \beta J(p,0)\right]\theta^{-1} + \frac{1}{2}\left(\left[U(\xi_j) + \beta J(p,0)\right]^2 - 2\beta\frac{\partial J(p;0)}{\partial \theta^{-1}}\right)\theta^{-2}.
\end{aligned}$$

Denote: $J_0(p) = J(p;0)$, $J_1(p) = \frac{\partial J(p;0)}{\partial \theta^{-1}}$, so that:

$$\begin{aligned}
\exp\left[-\frac{J(p;\theta^{-1})}{\theta}\right] &\approx 1 - J_0(p)\theta^{-1} + \frac{1}{2}\left[J_0(p)^2 - 2\frac{\partial J_0(p)}{\partial \theta^{-1}}\right]\theta^{-2} \\
\exp\left[-\frac{U(\xi_j) + \beta J(p;\theta^{-1})}{\theta}\right] &\approx 1 - \left[U(\xi_j) + \beta J_0(p)\right]\theta^{-1} + \frac{1}{2}\left(\left[U(\xi_j) + \beta J_0(p)\right]^2 - 2\beta J_1(p)\right)\theta^{-2}.
\end{aligned}$$

Substituting into the risk-sensitive recursion and equating coefficients on powers of θ^{-1}, we obtain the following system of equations for $J_0(p)$ and $J_1(p)$:

$$
\begin{aligned}
J_0(p) &= \sum_{j=1}^{N} p_j \left\{ \sum_{k\in a_B^c} \lambda_{jk} \, |2\pi\Sigma_u|^{-\frac{1}{2}} \int [U(\xi_j) + \beta J_0(p^*)] \, df(y^* - \xi_k) + \sum_{i=1}^{I} \lambda_{ji^*} \left[U(\xi_j) + \beta J_0^{\tau_i}\right] \right\} \\
2J_1(p) &= J_0(p)^2 + \sum_{j=1}^{N} p_j \sum_{k\in a_B^c} \lambda_{jk} \, |2\pi\Sigma_u|^{-\frac{1}{2}} \int \left(2\beta J_1(p^*) - [U(\xi_j) + \beta J_0(p^*)]^2\right) df(y^* - \xi_k) \\
&\quad + \sum_{j=1}^{N} p_j \sum_{i=1}^{I} \lambda_{ji^*} \left(2\beta J_1^{\tau_i} - [U(\xi_j) + \beta J_0^{\tau_i}]^2\right) \\
J_0^{\tau_i} &= \sum_{k\in a_B^c} \lambda_{i^*k} \, |2\pi\Sigma_u|^{-\frac{1}{2}} \int [U(\xi_{i^*}) + \beta J_0(p^*)] \, df(y^* - \xi_k) + \sum_{j=1}^{I} \lambda_{i^*j^*} \left[U(\xi_{i^*}) + \beta J_0^{\tau_j}\right] \\
2J_1^{\tau_i} &= (J_0^{\tau_i})^2 + \sum_{k\in a_B^c} \lambda_{i^*k} \, |2\pi\Sigma_u|^{-\frac{1}{2}} \int \left(2\beta J_1(p^*) - [U(\xi_{i^*}) + \beta J_0(p^*)]^2\right) df(y^* - \xi_k) \\
&\quad + \sum_{j=1}^{I} \lambda_{j^*i^*} \left(2\beta J_1^{\tau_j} - \left[U(\xi_{i^*}) + \beta J_0^{\tau_j}\right]^2\right).
\end{aligned}
$$

To solve the above system, notice that, since the unnormalized probability vector π is proportional to the conditional probability vector p, we can use π as the state variable as long as we recognize that J_0 and J_1 are homogeneous of degree 0 in π, so that $J_0(\alpha\pi) = J_0(\pi)$ and $J_1(\alpha\pi) = J_1(\pi)$ $\forall\alpha \in \mathbb{R}$. Recall that $p_j = \pi_j/(\mathbf{1}_N'\pi_j)$, so that, in terms of π, we can rewrite the above system as:

$$
\begin{aligned}
J_0(p) &= \sum_{j=1}^{N} \frac{\pi_j}{\mathbf{1}_N'\pi_j} \left\{ \sum_{k\in a_B^c} \lambda_{jk} \, |2\pi\Sigma_u|^{-\frac{1}{2}} \int [U(\xi_j) + \beta J_0(p^*)] \, df(y^* - \xi_k) + \sum_{i=1}^{I} \lambda_{ji^*} \left[U(\xi_j) + \beta J_0^{\tau_i}\right] \right\} \\
2J_1(p) &= J_0(p)^2 + \sum_{j=1}^{N} \frac{\pi_j}{\mathbf{1}_N'\pi_j} \sum_{k\in a_B^c} \lambda_{jk} \, |2\pi\Sigma_u|^{-\frac{1}{2}} \int \left(2\beta J_1(p^*) - [U(\xi_j) + \beta J_0(p^*)]^2\right) df(y^* - \xi_k) \\
&\quad + \sum_{j=1}^{N} \frac{\pi_j}{\mathbf{1}_N'\pi_j} \sum_{i=1}^{I} \lambda_{ji^*} \left(2\beta J_1^{\tau_i} - [U(\xi_j) + \beta J_0^{\tau_i}]^2\right)
\end{aligned}
$$

I look for a first order approximation to the solution in terms of log deviations from the steady state conditional distribution. In particular, denote by $\overline{\pi}$ the stationary distribution of the Markov chain $\{z_t\}_{t\geq 0}$: $\overline{\pi} = \Lambda'\overline{\pi}$. Notice that $\overline{\pi}$ is the steady state conditional distribution in the limiting case of arbitrarily uninformative signals where $\Sigma_u^{-1} = 0$. Introduce $\hat{\pi}$ to be the vector of log deviations from the stationary distribution:

$$
\text{(C.1)} \qquad \hat{\pi}_{jt} = \begin{cases} \log\tilde{\pi}_{jt} - \log\overline{\pi}_j; & j \in a_B^c \\ 0 & j \in a_B \end{cases},
$$

where $\tilde{\pi}_{jt} = f(y_t - \xi_j) \sum_{k=1}^{N} \lambda_{kj} \pi_{k,t-1}$ is the unnormalized probability vector before conditioning on observations of default at date t, and approximate:

$$\begin{aligned} J_0(\pi) &\approx J_0(\overline{\pi}) + \frac{\partial J_0(\overline{\pi})'}{\partial \pi} diag(\overline{\pi}) \hat{\pi} \\ J_0(\pi)^2 &\approx J_0(\overline{\pi})^2 + 2 J_0(\overline{\pi}) \frac{\partial J_0(\overline{\pi})'}{\partial \pi} diag(\overline{\pi}) \hat{\pi} \\ J_1(\pi) &\approx J_1(\overline{\pi}) + \frac{\partial J_1(\overline{\pi})'}{\partial \pi} diag(\overline{\pi}) \hat{\pi}. \end{aligned}$$

For simplicity, denote: $\gamma_{00} = J_0(\overline{\pi})$, $\gamma_{01} = \frac{\partial J_0(\overline{\pi})}{\partial \pi}$, $\gamma_{10} = J_1(\overline{\pi})$, and $\gamma_{11} = \frac{\partial J_1(\overline{\pi})}{\partial \pi}$, so that:

$$\begin{aligned} J_0(\pi) &\approx \gamma_{00} + \gamma_{01}' diag(\overline{\pi}) \hat{\pi} \\ J_0(\pi)^2 &\approx \gamma_{00}^2 + 2\gamma_{00} \gamma_{01}' diag(\overline{\pi}) \hat{\pi} \\ J_1(\pi) &\approx \gamma_{10} + \gamma_{11}' diag(\overline{\pi}) \hat{\pi}. \end{aligned}$$

Notice that the restrictions $J_0(\alpha\pi) = J_0(\pi)$ and $J_1(\alpha\pi) = J_1(\pi)$ imply that:

$$\begin{aligned} \sum_{j \in a_B^c} \gamma_{01,j} \overline{\pi}_j &= 0 \\ \sum_{j \in a_B^c} \gamma_{11,j} \overline{\pi}_j &= 0. \end{aligned}$$

Consider now the updating rule for log deviations from the steady state. Recall that unnormalized probabilities are updated according to:

$$\pi_i^* = f(y^* - \xi_j) \sum_{k=1}^{N} \lambda_{kj} \pi_k,$$

or, equivalently,

$$\frac{\pi_j^*}{\overline{\pi}_j} = f(y^* - \xi_j) \sum_{k \in a_B^c} \lambda_{kj} \frac{\overline{\pi}_k}{\overline{\pi}_j} \frac{\pi_k}{\overline{\pi}_k}.$$

Taking logs of both sides, we obtain the following updating rule for log deviations from the stationary distribution:

$$\hat{\pi}_j^* = \log f(y^* - \xi_j) + \log \left(\sum_{k \in a_B^c} \lambda_{kj} \frac{\overline{\pi}_k}{\overline{\pi}_j} \exp(\hat{\pi}_k) \right).$$

Approximating once again around $\hat{\pi} = \overrightarrow{0}$, we obtain:

$$\hat{\pi}_j^* \approx \log f(y^* - \xi_j) + \log \left(\sum_{k \in a_B^c} \lambda_{kj} \frac{\overline{\pi}_k}{\overline{\pi}_j} \right) + \left(\sum_{k \in a_B^c} \lambda_{kj} \overline{\pi}_k \right)^{-1} \sum_{k \in a_B^c} \lambda_{kj} \overline{\pi}_k \hat{\pi}_k.$$

Denote: $\mathcal{L}_{0j} = \log\left(\sum_{k\in a_B^c}\lambda_{kj}\frac{\overline{\pi}_k}{\overline{\pi}_j}\right)$, $\mathcal{L}_{1,jk} = \left(\sum_{k\in a_B^c}\lambda_{kj}\overline{\pi}_k\right)^{-1}\lambda_{kj}$, so that the first order approximation to the evolution equation is given by:

$$\hat{\pi}_j^* \approx \log f(y^* - \xi_k) + \mathcal{L}_{0j} + \mathcal{L}'_{1j} diag(\overline{\pi})\hat{\pi}.$$

Notice also that, after observing the default state of firm i, i^*, the predicted vector of conditional distribution is given by:

$$\hat{\pi}_j^* = \log f(y^* - \xi_j) + \log\lambda_{i^*j} - \log\overline{\pi}_j, \qquad \forall j \in a_B^c.$$

Similarly, approximate:

$$p_j = \frac{\pi_j}{\mathbf{1}'_N\pi} =\approx \overline{\pi}_j(1 + \hat{\pi}_j - \overline{\pi}'\hat{\pi}).$$

Substituting into the above system and equating coefficients, we obtain the following system:

$$\text{(C.2)} \qquad \gamma_{00} = \sum_{j\in a_B^c}\overline{\pi}_j\left\{U(\xi_j) + \beta\sum_{k\in a_B^c}\lambda_{jk}\left(\gamma_{00} + \gamma'_{01}diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right)\right\} + \beta\sum_{j\in a_B^c}\overline{\pi}_j\sum_{i=1}^{I}\lambda_{ji^*}J_0^{\tau_i}$$

$$\text{(C.3)} \qquad \gamma_{01,j} = U(\xi_j) + \beta\sum_{k\in a_B^c}\lambda_{jk}\left(\gamma_{00} + \gamma'_{01}diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right) + \beta\sum_{i=1}^{I}\lambda_{ji^*}J_0^{\tau_i} - \gamma_{00} + \beta\gamma'_{01}diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})\mathcal{L}'_{1j}\left(\sum_{l,k\in a_B^c}\overline{\pi}_l\lambda_{lk}\right)$$

$$\text{(C.4)} \qquad J_0^{\tau_i} = \sum_{k\in a_B^c}\lambda_{i^*k}\left(U(\xi_{i^*}) + \beta\left[\gamma_{00} + \gamma'_{01}diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})\left(\Delta_k^1 + \mathcal{L}_{0i^*}^d\right)\right]\right) + \sum_{j=1}^{I}\lambda_{i^*j^*}\left[U(\xi_{i^*}) + \beta J_0^{\tau_j}\right]$$

$$
\begin{aligned}
\text{(C.5)}\ 2\gamma_{10} &= \gamma_{00}^2 - \sum_{j\in a_B^c} \overline{\pi}_j U(\xi_j)^2 + 2\beta \sum_{j,k\in a_B^c} \overline{\pi}_j \lambda_{jk} \left(\gamma_{10} + \gamma_{11}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right) \\
&\quad + \beta \sum_{j\in a_B^c} \overline{\pi}_j \sum_{i=1}^{I} \lambda_{ji^*} \left[2J_1^{\tau_i} - 2U(\xi_j) J_0^{\tau_i} - \beta \left(J_0^{\tau_i}\right)^2\right] \\
&\quad - 2\beta \sum_{j,k\in a_B^c} \overline{\pi}_j \lambda_{jk} U(\xi_j) \left(\gamma_{00} + \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right) \\
&\quad + \beta^2 \gamma_{00} \sum_{j,k\in a_B^c} \overline{\pi}_j \lambda_{jk} U(\xi_j) \left(\gamma_{00} + 2\gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right) \\
\text{(C.6)}\ 2\gamma_{11,j} &= 2\gamma_{00}\gamma_{01,j} + \gamma_{00}^2 - 2\gamma_{10} - U(\xi_j)^2 \\
&\quad + 2\beta \sum_{k\in a_B^c} \lambda_{jk} \left(\gamma_{10} + \gamma_{11}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right) \\
&\quad + \beta \sum_{i=1}^{I} \lambda_{ji^*} \left[2J_1^{\tau_i} - 2U(\xi_j) J_0^{\tau_i} - \beta \left(J_0^{\tau_i}\right)^2\right] \\
&\quad - 2\beta \sum_{k\in a_B^c} \lambda_{jk} U(\xi_j) \left(\gamma_{00} + \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right) \\
&\quad + \beta^2 \gamma_{00} \sum_{k\in a_B^c} \lambda_{jk} U(\xi_j) \left(\gamma_{00} + 2\gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_0]\right) \\
&\quad + 2\beta \left(\gamma_{11} - (1-\beta\gamma_{00})\gamma_{01}\right)' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \mathcal{L}_{1j}' \left(\sum_{l,k\in a_B^c} \overline{\pi}_l \lambda_{lk}\right) \\
\text{(C.7)}\ 2J_1^{\tau_i} &= \left(J_0^{\tau_i}\right)^2 + 2\beta \sum_{k\in a_B^c} \left(\gamma_{10} + \gamma_{11}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_{0i}^d]\right) \\
&\quad + \sum_{j=1}^{I} \lambda_{i^*j^*} \left[2\beta J_1^{\tau_j} - \left(U(\xi_{i^*}) + \beta J_0^{\tau_j}\right)\right] - U(\xi_{i^*})^2 \sum_{k\in a_B^c} \lambda_{i^*k} \\
&\quad - 2\beta U(\xi_{i^*}) \sum_{k\in a_B^c} \lambda_{i^*k} \left(\gamma_{00} + \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_{0i}^d]\right) \\
&\quad + \beta^2 \gamma_{00} \sum_{k\in a_B^c} \lambda_{i^*k} \left(\gamma_{00} + 2\gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})[\Delta_k^1 + \mathcal{L}_{0i}^d]\right),
\end{aligned}
$$

where Δ_k^1 is a constant vector given by:

$$\Delta_{kj}^1 = -\frac{1}{2} - \frac{1}{2}(\xi_j - \xi_k)' \Sigma_u^{-1} (\xi_j - \xi_k).$$

Consider now the distortion to the conditional joint distribution of next period's signals and state. Recall that, in terms of the value function, this is given by:

$$\phi_t(z^*, y^*; \theta^{-1}) = \frac{\exp\left(-\frac{\beta J(\pi^*;\theta^{-1})}{\theta}\right)}{\mathbb{E}\left[\exp\left(-\frac{\beta J(\pi^*;\theta^{-1})}{\theta}\right)\Big| \mathcal{F}_t\right]}.$$

Taking a first order around the point $\theta^{-1} = 0$, we obtain:

$$\phi_0(z^*, y^*|z = \xi_j) + \theta^{-1}\phi_1(z^*, y^*|z = \xi_j) = 1 - \theta^{-1}\beta\left(J_0(\pi^*) - \mathbb{E}\left[J_0(\pi^*)|\, z_t = \xi_j\right]\right).$$

Equating coefficients, we obtain:

$$\begin{aligned}\phi_0(z^*, y^*|z = \xi_j) &= 1\\ \phi_1(z^*, y^*|z = \xi_j) &= -\beta\left(J_0(\pi^*) - \mathbb{E}\left[J_0(\pi^*)|\, z_t = \xi_j\right]\right).\end{aligned}$$

Substituting for J_0, we obtain:

$$\phi_1(z^*, y^*|z = \xi_j) = \varphi_0 + \varphi_\pi' diag(\overline{\pi})\hat{\pi} + \varphi_y' \log f(y^*),$$

where:

$$\text{(C.8)}\ \varphi_{0,jk} = \begin{cases} \beta\left(\sum_{i=1}^I \lambda_{ji^*} J_0^{\tau_i} - \gamma_{00} - \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\mathcal{L}_0\right) \\ +\beta\sum_{l\in a_B^c}\lambda_{jl}\left[\gamma_{00} + \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\Delta_k^1 + \mathcal{L}_0)\right] & j,k\in a_B^c \\ \beta\left(\sum_{i=1}^I \lambda_{ji^*} J_0^{\tau_i} - J_0^{\tau_k}\right) + \\ +\beta\sum_{l\in a_B^c}\lambda_{jl}\left[\gamma_{00} + \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\Delta_k^1 + \mathcal{L}_0)\right] & j\in a_B^c, k\in a_B \\ \beta\left(\sum_{i=1}^I \lambda_{j^*i^*} J_0^{\tau_i} - \gamma_{00} - \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\mathcal{L}_{0j}^d\right) \\ +\beta\sum_{l\in a_B^c}\lambda_{j^*l}\left[\gamma_{00} + \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\mathcal{L}_{0j}^d\right] & j\in a_B, k\in a_B^c \\ \beta\left(\sum_{i=1}^I \lambda_{j^*i^*} J_0^{\tau_i} - \gamma_{00} - \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\mathcal{L}_{0j}^d\right) \\ +\beta\sum_{l\in a_B^c}\lambda_{j^*l}\left[\gamma_{00} + \gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\mathcal{L}_{0j}^d\right] & j,k\in a_B \end{cases}$$

$$\text{(C.9)}\ \varphi_{\pi,jk} = \begin{cases} \beta\gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\mathcal{L}_1\left(\sum_{l\in a_B^c}\lambda_{jl} - 1\right) & j,k\in a_B^c \\ \beta\mathcal{L}_1' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\gamma_{01}\sum_{k\in a_B^c}\lambda_{jk} & j\in a_B^c, k\in a_B \\ 0 & j\in a_B \end{cases}$$

$$\text{(C.10)}\ \varphi_{y,jk} = \begin{cases} -\beta\gamma_{01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) & j,k\in a_B^c \\ -\beta diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})\gamma_{01} & j\in a_B, k\in a_B^c \\ 0 & k\in a_B \end{cases}$$

Turn now to the distortion to the current period's conditional probability vector. Recall that this is given by:

$$\psi_t(z) = \frac{\exp\left(-\frac{U(z)+\mathcal{R}_t(\beta J(p^*))}{\theta}\right)}{\mathbb{E}\left[\exp\left(-\frac{U(z)+\mathcal{R}_t(\beta J(p^*))}{\theta}\right)\Big| \mathcal{G}_t\right]}.$$

Approximating once again around the point $\theta^{-1} = 0$, we obtain:

$$\begin{aligned}\psi_0(\xi_j) + \theta^{-1}\psi_1(\xi_j) &= 1 - \theta^{-1}\sum_{k=1}^{N}\lambda_{jk}\,|2\pi\Sigma_u|^{-\frac{1}{2}}\int\left[U(\xi_j) + \beta J_0(\pi^*)\right]df(y^* - \xi_k) \\ &\quad +\theta^{-1}\sum_{k,l=1}^{N}\lambda_{lk}\,|2\pi\Sigma_u|^{-\frac{1}{2}}\int\left[U(\xi_l) + \beta J_0(\pi^*)\right]df(y^* - \xi_k).\end{aligned}$$

Substituting for $J_0(\pi^*)$ and equating coefficients, we obtain:

$$\begin{aligned}\psi_0(\xi_j) &= 1 \\ \psi_1(\xi_j) &= \sum_{l=1}^{N} p_l\left[U(\xi_l) - U(\xi_j)\right] + \beta\sum_{l=1}^{N} p_l\sum_{i=1}^{I}(\lambda_{li^*} - \lambda_{ji^*})J_0^{\tau_i} \\ &\quad +\beta\sum_{l=1}^{N} p_l\sum_{k\in a_B^c}(\lambda_{lk} - \lambda_{jk})\left[\gamma_{00} + \gamma_{01}'diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})\left[\Delta_k^1 + \mathcal{L}_0 + \mathcal{L}_1 diag(\overline{\pi})\hat{\pi}\right]\right]\end{aligned}$$

Recall that, in terms of log-deviations from the steady state, $p_l \approx \overline{\pi}_l\,(1 + \hat{\pi}_l - \overline{\pi}'\hat{\pi})$. We can represent:

$$\psi_1(\xi_j) = \zeta_{j0} + \zeta_{j1}'diag(\overline{\pi})\hat{\pi},$$

where:

$$\begin{aligned}\text{(C.11)}\qquad \zeta_{j0} &= \sum_{l\in a_B^c}\overline{\pi}_l\left\{U(\xi_l) + \beta\sum_{k\in a_B^c}\lambda_{lk}\left[\gamma_{00} + \gamma_{01}'diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})(\Delta_k^1 + \mathcal{L}_0)\right]\right\} \\ &\quad +\beta\sum_{l\in a_B^c}\overline{\pi}_l\sum_{i=1}^{I}\lambda_{li^*}J_0^{\tau_k} - U(\xi_l) \\ &\quad -\beta\sum_{k\in a_B^c}\lambda_{lk}\left[\gamma_{00} + \gamma_{01}'diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})(\Delta_k^1 + \mathcal{L}_0)\right] - \beta\sum_{i=1}^{I}\lambda_{li^*}J_0^{\tau_k} \\ \text{(C.12)}\qquad \zeta_{j1,k} &= -\zeta_{k0} + \beta\gamma_{01}'diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})\mathcal{L}_{1k}'\left(\sum_{l,m\in a_B^c}\overline{\pi}_l\lambda_{lm} - \sum_{l\in a_B^c}\lambda_{kl}\right).\end{aligned}$$

C.2 Proof of Lemma 3.3

Notice first that, substituting the first order expansion to the value of equity of firm i into the worst-case Euler equation and equating coefficients on the powers of θ^{-1}, we obtain the following system for V_{i0} and V_{i1}:

$$\begin{aligned}
V_{i0}(\pi) &= -C_i + \sum_{j \in a_B^c} \left[\delta_i A_{ij} + \beta \sum_{k \in a_B^c} \lambda_{jk} \left| 2\pi\Sigma_u \right|^{-\frac{1}{2}} \int V_{i0}(\pi^*) df(y^* - \xi_k) + \beta \sum_{k \neq i} \lambda_{jk^*} V_{i0}^{\tau_k} \right] \\
V_{i0}^{\tau_j} &= -C_i + \delta_i A_{ij^*} + \beta \left[\sum_{k \in a_B^c} \lambda_{j^*k} \left| 2\pi\Sigma_u \right|^{-\frac{1}{2}} \int V_{i0}(\pi^*) df(y^* - \xi_k) + \sum_{k \neq i} \lambda_{j^*k^*} V_{i0}^{\tau_k} \right]
\end{aligned}$$

$$\begin{aligned}
V_{i1}(\pi) &= \sum_{j \in a_B^c} p_j \psi_1(\xi_j) \left[\delta_i A_{ij} + \beta \sum_{k \in a_B^c} \lambda_{jk} \left| 2\pi\Sigma_u \right|^{-\frac{1}{2}} \int V_{i0}(\pi^*) df(y^* - \xi_k) + \beta \sum_{k \neq i} \lambda_{jk^*} V_{i0}^{\tau_k} \right] \\
&\quad + \beta \sum_{j,k \in a_B^c} p_j \lambda_{jk} \left| 2\pi\Sigma_u \right|^{-\frac{1}{2}} \int \left[V_{i0}(\pi^*) \phi_1(\xi_k, y^* | \xi_j) + V_{i1}(\pi^*) \right] df(y^* - \xi_k) \\
&\quad + \beta \sum_{j \in a_B^c} p_j \sum_{k \neq i} \lambda_{jk^*} \left[V_{i0}^{\tau_k} \phi_1(\xi_{k^*} | \xi_j) + V_{i1}^{\tau_k} \right] \\
V_{i1}^{\tau_j} &= \beta \sum_{k \in a_B^c} \lambda_{j^*k} \left| 2\pi\Sigma_u \right|^{-\frac{1}{2}} \int \left[V_{i0}(\pi^*) \phi_1(\xi_k, y^* | \xi_j) + V_{i1}(\pi^*) \right] df(y^* - \xi_k) \\
&\quad + \beta \sum_{k \neq i} \lambda_{j^*k^*} \left[V_{i0}^{\tau_k} \phi_1(\xi_{k^*} | \xi_j^*) + V_{i1}^{\tau_k} \right],
\end{aligned}$$

where A_{ij} is the fundamental value of asset of firm i in state j. Substituting the approximations (3.22)-(3.23) and equating coefficients, we obtain:

$$\begin{aligned}
\nu_{i,00} &= -C_i + \sum_{j \in a_B^c} \overline{\pi}_j \left(\delta_i A_{ij} + \beta \sum_{k \in a_B^c} \lambda_{jk} \left[\nu_{i,00} + \nu_{i,01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) (\mathcal{L}_0 + \Delta_k^1) \right] + \beta \sum_{k \neq i} \lambda_{jk^*} V_{i0}^{\tau_k} \right) \\
\nu_{i,01,j} &= \delta_i A_{ij} + \beta \sum_{k \in a_B^c} \lambda_{jk} \left[\nu_{i,00} + \nu_{i,01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) (\mathcal{L}_0 + \Delta_k^1) \right] + \beta \sum_{k \neq i} \lambda_{jk^*} V_{i0}^{\tau_k} \\
&\quad - \nu_{i,00} + \beta \sum_{k \neq i} \lambda_{jk^*} V_{i0}^{\tau_k} + \beta \nu_{i,01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \mathcal{L}_{1j}' \sum_{l,k \in a_B^c} \overline{\pi}_l \lambda_{lk} \\
V_{i0}^{\tau_j} &= -C_i + \delta_i A_{ij^*} + \beta \sum_{k \neq i} \lambda_{j^*k^*} V_{i0}^{\tau_k} + \beta \sum_{k \in a_B^c} \lambda_{j^*k} \left[\nu_{i,00} + \nu_{i,01}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \left(\Delta_k^1 + \mathcal{L}_{0j}^d \right) \right]
\end{aligned}$$

$$
\begin{aligned}
\nu_{i,10} &= \sum_{j\in a_B^c} \overline{\pi}_j \zeta_{j0} \left\{ \delta_i A_{ij} + \beta \sum_{k\neq i} \lambda_{jk^*} V_{i0}^{\tau_k} + \beta \sum_{k\in a_B^c} \lambda_{jk} \left[\nu_{i,00} + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right] \right\} \\
&+ \beta \sum_{j,k\in a_B^c} \overline{\pi}_j \lambda_{jk} \left[\nu_{i,10} + \nu'_{i,11} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right] \\
&+ \beta \sum_{j,k\in a_B^c} \overline{\pi}_j \lambda_{jk} \left[\nu_{i,00} + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right] \varphi_{0,jk} \\
&+ \beta \sum_{j,k\in a_B^c} \overline{\pi}_j \lambda_{jk} \left[\nu_{i,00} \varphi'_{y,jk} \Delta_k^1 + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \Delta_k^2 \varphi_{y,jk} \right] \\
&+ \beta \sum_{j\in a_B^c} \overline{\pi}_j \sum_{k\neq i} \lambda_{jk^*} \left[V_{i0}^{\tau_k} \varphi_{0,jk^*} + V_{i1}^{\tau_k} \right]
\end{aligned}
$$

$$
\begin{aligned}
\nu_{i,11,j} &= \zeta_{j0} \left\{ \delta_i A_{ij} + \beta \sum_{k\neq i} \lambda_{jk^*} V_{i0}^{\tau_k} + \beta \sum_{k\in a_B^c} \lambda_{jk} \left[\nu_{i,00} + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right] \right\} \\
&+ \beta \sum_{k\in a_B^c} \lambda_{jk} \left[\nu_{i,10} + \nu'_{i,11} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right] \\
&+ \beta \sum_{k\in a_B^c} \lambda_{jk} \left[\nu_{i,00} + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right] \varphi_{0,jk} - \nu_{i,10} \\
&+ \beta \sum_{k\in a_B^c} \lambda_{jk} \left[\nu_{i,00} \varphi'_{y,jk} \Delta_k^1 + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \Delta_k^2 \varphi_{y,jk} \right] + \beta \sum_{k\neq i} \lambda_{jk^*} \left[V_{i0}^{\tau_k} \varphi_{0,jk^*} + V_{i1}^{\tau_k} \right] \\
&+ \sum_{l\in a_B^c} \zeta_{l1,j} \left[\delta_i A_{il} + \beta \sum_{k\neq i} \lambda_{lk^*} V_{i0}^{\tau_k} + \beta \sum_{k\in a_B^c} \lambda_{lk} \left(\nu_{i,00} + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right) \right] \\
&+ \beta \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \mathcal{L}'_{1j} \sum_{l,k\in a_B^c} \overline{\pi}_l \zeta_{l0} \lambda_{lk} + \beta \nu'_{i,11} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \mathcal{L}'_{1j} \sum_{l,k\in a_B^c} \overline{\pi}_l \zeta_{l0} \lambda_{lk} \\
&+ \beta \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \mathcal{L}'_{1j} \sum_{l,k\in a_B^c} \overline{\pi}_l \lambda_{lk} \left[\varphi_{0,lk} + \varphi'_{y,lk} \Delta_k^1 \right] + \beta \sum_{l\in a_B^c} \sum_{k\neq i} \overline{\pi}_l \lambda_{lk^*} V_{i0}^{\tau_k} \varphi_{\pi,lk^*,j} \\
&+ \beta \sum_{l,k\in a_B^c} \overline{\pi}_l \lambda_{lk} \left[\nu_{i,00} + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_0 + \Delta_k^1) \right] \varphi_{\pi,lk,j} \\
V_{i1}^{\tau_j} &= \beta \sum_{k\in a_B^c} \lambda_{j^*k} \left[\nu_{i,00} + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_{0j}^d + \Delta_k^1) \right] \varphi_{0,j^*k} \\
&+ \beta \sum_{k\in a_B^c} \lambda_{j^*k} \left[\nu_{i,00} \varphi'_{y,j^*k} \Delta_k^1 + \nu'_{i,01} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \Delta_k^2 \varphi_{y,j^*k} \right] \\
&+ \beta \sum_{k\in a_B^c} \lambda_{j^*k} \left[\nu_{i,10} + \nu'_{i,11} diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c})(\mathcal{L}_{0j}^d + \Delta_k^1) \right] + \beta \sum_{k\neq i} \lambda_{j^*k^*} \left(V_{i0}^{\tau_k} \phi_1(\xi_{k^*}|\xi_{j^*}) + V_{i1}^{\tau_k} \right)
\end{aligned}
$$

where Δ_k^2 is a constant matrix given by:

$$\begin{aligned}\Delta^2_{k,jl} &\equiv |2\pi\Sigma_u|^{-\frac{1}{2}} \int \left(\frac{1}{2}(y^* - \xi_j - \overline{u})'\Sigma_u^{-1}(y^* - \xi_j - \overline{u})\right)\left(\frac{1}{2}(y^* - \xi_j - \overline{u})'\Sigma_u^{-1}(y^* - \xi_j - \overline{u})\right) df(y^* - \xi_k) \\ &= \frac{1}{4}\left[(\xi_k - \xi_j)'\Sigma_u^{-1}(\xi_k - \xi_j)\right]\left[(\xi_k - \xi_l)'\Sigma_u^{-1}(\xi_k - \xi_l)\right] + \frac{1}{4}(\xi_k - \xi_j)'\Sigma_u^{-1}(\xi_k - \xi_j) \\ &\quad + \frac{1}{4}(\xi_k - \xi_l)'\Sigma_u^{-1}(\xi_k - \xi_l) + (\xi_k - \xi_l)'\Sigma_u^{-1}(\xi_k - \xi_j) + \frac{5}{4}.\end{aligned}$$

D Risk aversion benchmark

In this section, I investigate the performance of a power utility model in explaining the time series evolution of credit spreads and equity prices. Instead of solving the portfolio allocation problem of the representative risk-averse agent, I take the stochastic discount factor as given. In particular, let $S_{t,t+s}$ be the stochastic discount factor between dates t and $t+s$. In an economy where the representative agent evaluates consumption paths using a power utility function:

$$u(\mathcal{K}_t) = \frac{\mathcal{K}_t^{1-\gamma}}{1-\gamma},$$

where $\mathcal{K}_t \equiv \mathcal{K}(z_t) = \sum_{i=1}^{I} \delta_i A_i(z_t)$ is the level of consumption at date t and $\gamma > 0$ is the degree of risk aversion, the stochastic discount factor is given by:

$$S_{t,t+s} = \beta^s \frac{\mathcal{K}_{t+s}^{-\gamma}}{\mathcal{K}_t^{-\gamma}}. \tag{D.1}$$

Consider first the CDS spread on a swap with maturity $T = t + 6n$ on the consol bond of firm i. In the economy with the risk-averse representative agent, this is given by:

$$c_i(t,T) = \frac{2X_i \mathbb{E}\left[\beta^{\tau_i - t}\mathcal{K}_{\tau_i}^{-\gamma}\mathbf{1}_{\tau_i < T}\middle|\, \mathcal{G}_t\right]}{\sum_{s=1}^{n}\beta^{6s}\mathbb{E}\left[\mathcal{K}_{t+6s}^{-\gamma}\mathbf{1}_{\tau_i \geq t+6s}\middle|\, \mathcal{G}_t\right]}.$$

Notice that:

$$\begin{aligned}\mathbb{E}\left[\beta^{\tau_i - t}\mathcal{K}_{\tau_i}^{-\gamma}\mathbf{1}_{\tau_i < T}\middle|\, \mathcal{G}_t\right] &= \sum_{s=1}^{T-t}\beta^s \mathbb{E}\left[\mathbf{1}_{\tau_i = t+s}\mathcal{K}_{t+s}^{-\gamma}\middle|\, \mathcal{G}_t\right] \\ \mathbb{E}\left[\mathcal{K}_{t+6s}^{-\gamma}\mathbf{1}_{\tau_i \geq t+6s}\middle|\, \mathcal{G}_t\right] &= \mathbb{E}\left[\mathcal{K}_{t+6s}^{-\gamma}\middle|\, \mathcal{G}_t\right] - \sum_{k=1}^{6s}\mathbb{E}\left[\mathcal{K}_{t+6s}^{-\gamma}\mathbf{1}_{\tau_i = t+k}\middle|\, \mathcal{G}_t\right].\end{aligned}$$

Let $q_{ij}^{(n)} = \mathbb{P}\left(z_{t+n} = \xi_j,\ z_{t+n-1} \neq \xi_j, \ldots,\ z_{t+1} \neq \xi_j \,|\, z_t = \xi_i\right)$ be the probability that $t+n$ is the first hitting time of ξ_j conditional on being in state i at date t. Then:

$$\begin{aligned}
\mathbb{E}\left[\mathcal{K}_{t+s}^{-\gamma}\mathbf{1}_{\tau_i=t+s}\middle|\,\mathcal{G}_t\right] &= \sum_{j=1}^{N} p_{jt}q_{ji^*}^{(s)}\mathcal{K}(a_{B_i})^{-\gamma} \\
\mathbb{E}\left[\mathcal{K}_{t+s}^{-\gamma}\middle|\,\mathcal{G}_t\right] &= \sum_{j\in a_B^c} p_{jt}\sum_{k=1}^{N}\{\Lambda^s\}_{jk}\,\mathcal{K}(\xi_k)^{-\gamma} \\
\mathbb{E}\left[\mathcal{K}_{t+s_1}^{-\gamma}\mathbf{1}_{\tau_i=t+s_2}\middle|\,\mathcal{G}_t\right] &= \sum_{j\in a_B^c} p_{jt}q_{ji^*}\sum_{k=1}^{N}\left\{\Lambda^{s_1-s_2}\right\}_{jk}\mathcal{K}(\xi_k)^{-\gamma}.
\end{aligned}$$

Table 7 presents the 5 year CDS spreads on the financial institutions for different levels of risk aversion at four of interest – before the start of the crisis, July 2007, at the start of the crisis in August 2007, after the bailout of Bear Stearns in March 2008 and after the liquidation of Lehman Brothers in September 2008 – together with the observations of the 5 year CDS spreads at these dates. Notice first that while the model-implied credit spreads do increase during the crisis, the magnitude of the model-implied spreads remains much smaller than that of the observed CDS spreads. The only exception is JP Morgan, the model-implied spreads for which exceed the observed ones during the crisis. Notice also, that the dependence on the degree of risk aversion is not monotone. In particular, although the spreads increase initially as the degree of risk aversion increases, further increases in risk aversion decrease the model-implied CDS spreads.

Consider now the date t price of a claim to the equity of firm i. Using the stochastic discount factor of the risk-averse agent, the equity price satisfies the Euler equation:

$$V_{it} = \mathbb{E}\left[\delta_i A_{it} - C_i + S_{t,t+1}\mathbf{1}_{\tau_i>t+1}V_{i,t+1}\middle|\,\mathcal{G}_t\right]. \tag{D.2}$$

Similarly to the case with model misspecification, I look for a first order approximation to the equity price in terms of log deviations from the steady state:

$$V_{ir}(\pi_t) = \nu_{ir,0} + \nu_{ir,1}' diag(\overline{\pi})\hat{\pi}_t + O_2(\hat{\pi}_t).$$

Substituting into the Euler equation and equation coefficients, we obtain:

$$\begin{aligned}
\nu_{ir,0} &= \sum_{j\in a_B^c}\overline{\pi}_j\left\{\delta_i A_{ij} - C_i + \beta\sum_{k\in a_B^c}\frac{\mathcal{K}(\xi_k)^{-\gamma}}{\mathcal{K}(\xi_j)^{-\gamma}}\lambda_{jk}\left[\nu_{ir,0} + \nu_{ir,1}' diag(\overline{\pi})diag(\mathbf{1}_{a_B^c}\left(\mathcal{L}_0 + \Delta_k^1\right)\right]\right\} \\
&\quad + \beta\sum_{j\in a_B^c}\overline{\pi}_j\sum_{k\neq i}\lambda_{jk^*}\frac{\mathcal{K}(\xi_{k^*})^{-\gamma}}{\mathcal{K}(\xi_j)^{-\gamma}} \\
\nu_{ir,1j} &= \delta_i A_{ij} - C_i + \beta\sum_{k\in a_B^c}\frac{\mathcal{K}(\xi_k)^{-\gamma}}{\mathcal{K}(\xi_j)^{-\gamma}}\lambda_{jk}\left[\nu_{ir,0} + \nu_{ir,1}' diag(\overline{\pi})diag(\mathbf{1}_{a_B^c}\left(\mathcal{L}_0 + \Delta_k^1\right)\right] \\
&\quad + \beta\sum_{k\neq i}\lambda_{jk^*}\frac{\mathcal{K}(\xi_{k^*})^{-\gamma}}{\mathcal{K}(\xi_j)^{-\gamma}} - \nu_{ir,0} + \beta\nu_{ir,1}' diag(\overline{\pi})diag(\mathbf{1}_{a_B^c})\mathcal{L}_{1j}'\sum_{l,k\in a_B^c}\overline{\pi}_l\lambda_{lk}\frac{\mathcal{K}(\xi_k)^{-\gamma}}{\mathcal{K}(\xi_l)^{-\gamma}}.
\end{aligned}$$

Similarly, at the time of default of firm j, the equity price of firm i solves:

$$
\begin{aligned}
V_{ir}^{\tau_j} &= \delta_i A_{ij^*} - C_i + \beta \sum_{k \in a_B^c} \frac{\mathcal{K}(\xi_k)^{-\gamma}}{\mathcal{K}(\xi_{j^*})^{-\gamma}} \lambda_{j^* k} \left[\nu_{ir,0} + \nu_{ir,1}' diag(\overline{\pi}) diag(\mathbf{1}_{a_B^c}) \left(\mathcal{L}_0 + \mathcal{L}_1 diag(\overline{\pi}) \hat{\pi} + \Delta_k^1 \right) \right] \\
&\quad + \beta \sum_{k \neq i} \lambda_{j^* k^*} \frac{\mathcal{K}(\xi_{k^*})^{-\gamma}}{\mathcal{K}(\xi_{j^*})^{-\gamma}} V_{ir}^{\tau_k}.
\end{aligned}
$$

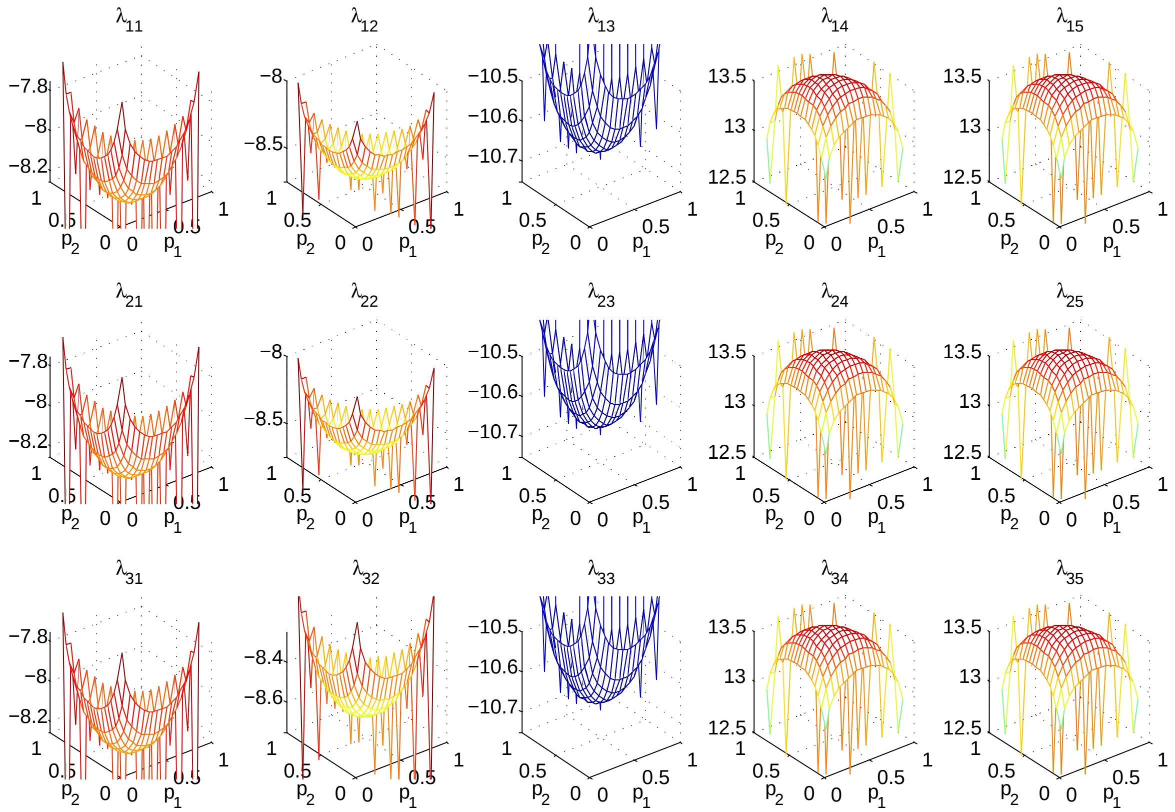

Figure 1: Differences between the reference transition probabilities and the misspecification-implied transition probabilities for $\theta = 0.5$. The parameters used are those of Section 3.3.1.

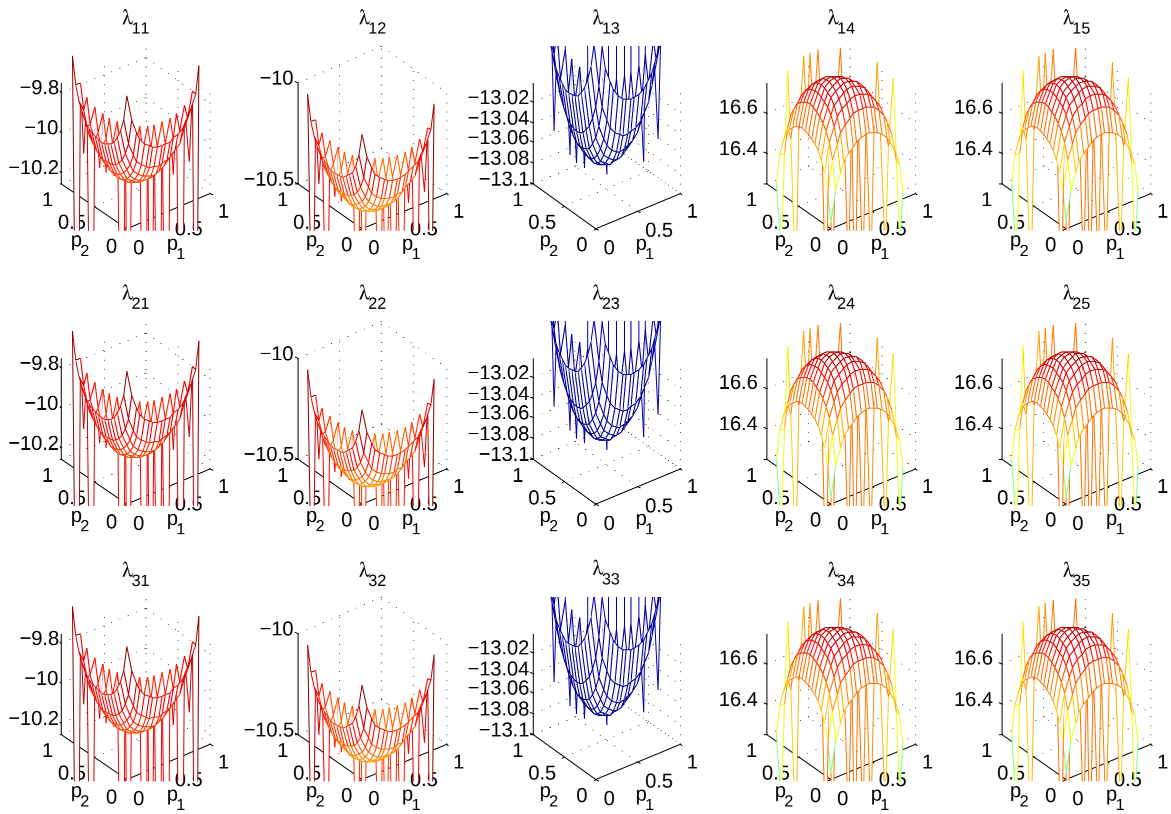

Figure 2: Differences between the reference transition probabilities and the misspecification-implied transition probabilities for $\theta = 1$. The parameters used are those of Section 3.3.1.

	Mean	St. Dev.	Correlations										
			BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC	CS10	SP500
BAC	56038.09	65351.99	100.00									76.99	61.27
BSC	4508.48	3687.57	98.08	100.00								95.85	83.32
Citi	53210.59	51820.79	91.39	56.58	100.00							91.01	75.66
GS	32395.77	18845.19	96.01	-20.99	93.12	100.00						48.25	-10.57
JPM	51611.38	57647.31	79.25	33.03	74.92	78.50	100.00					67.35	48.97
MER	15995.35	13145.67	88.73	73.51	97.45	92.10	71.71	100.00				95.81	76.41
MS	16683.24	14747.40	95.92	43.47	97.14	96.66	76.59	78.11	100.00			83.31	73.77
WB	21541.91	22530.87	96.66	92.99	93.97	96.84	71.40	95.30	93.99	100.00		93.78	75.48
WFC	23943.94	28543.76	96.35	19.57	88.91	94.29	73.87	63.31	94.59	31.24	100.00	68.15	55.87

Table 1: Sample moments of and empirical correlations between the book equity of financial institutions and aggregate market indicators. CS10 is the Case-Schiller 10 index; SP500 is the Standard and Poor's 500 index. Data Source: Compustat, CRSP

	BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC
δ	3.67	3.41	3.97	4.46	9.67	3.39	4.69	3.13	3.13
D	1398608	381237.992	2093112	891128	1338831	1034133	1160482	647525	492564
X	47.61	70.20	67.91	98.77	75.05	66.96	65.11	45.07	46.35
C	530.96	120.85	767.83	669.21	496.40	305.24	404.80	225.59	202.27

Table 2: Parameters estimated outside the MCMC procedure of Appendix A. δ, the fraction of assets generated as payoffs, and X, the payment in case of default, are reported in percentage terms; D is the face value of the consol bond, taken to be the last available observation of long-term debt; C is the monthly coupon payment on the consol bond.

ζ_f

BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC
7.40	7.98	9.66	5.54	10.00	9.21	9.11	9.37	9.13
9.83	6.92	9.58	5.29	9.99	9.10	9.00	9.28	9.02
9.75	7.93	6.91	5.19	9.92	9.08	8.98	9.27	8.99
9.74	7.91	9.55	0.13	9.91	9.04	8.97	9.25	8.95
9.86	7.99	9.71	5.57	8.00	9.01	8.91	9.17	8.86
9.98	8.07	9.90	5.99	10.11	8.02	8.87	9.16	8.85
9.89	7.99	9.73	5.66	10.03	9.12	6.82	9.14	8.85
9.73	7.91	9.54	5.19	9.91	8.98	8.88	7.49	8.84
9.75	7.92	9.59	5.21	9.95	9.01	8.90	9.19	7.05
9.73	7.88	9.52	5.05	9.90	8.99	8.87	9.14	8.84
9.74	7.94	9.58	5.28	9.91	9.04	8.91	9.16	8.88
9.86	8.01	9.73	5.66	10.02	9.12	8.99	9.27	8.98
9.69	7.89	9.49	5.10	9.90	8.98	8.85	9.10	8.80
9.71	7.89	9.54	5.16	9.88	9.02	8.87	9.13	8.85
9.71	7.91	9.55	5.14	9.90	9.01	8.87	9.13	8.83
9.78	7.94	9.60	5.37	9.95	9.03	8.93	9.18	8.89
9.75	7.92	9.57	5.30	9.93	9.03	8.92	9.18	8.87
9.83	7.97	9.67	5.47	10.00	9.09	8.98	9.23	8.96

ζ_c

4.06	4.82

ρ

BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC
0.20	0.00	0.09	0.01	0.27	0.09	0.09	0.06	0.01

Ω_c

99.32	0.68
0.54	99.46

Ω_f

54.90	2.74	2.63	2.56	2.71	2.68	2.60	2.58	2.66	2.70	2.53	2.63	2.73	2.64	2.63	2.63	2.70	2.75
2.58	55.04	2.61	2.68	2.69	2.73	2.66	2.66	2.65	2.66	2.59	2.57	2.62	2.64	2.69	2.59	2.74	2.61
2.63	2.65	55.12	2.65	2.60	2.59	2.72	2.66	2.71	2.68	2.67	2.65	2.59	2.69	2.57	2.57	2.63	2.62
2.63	2.65	2.61	55.08	2.63	2.75	2.53	2.63	2.68	2.62	2.65	2.65	2.63	2.69	2.67	2.62	2.69	2.62
2.62	2.64	2.57	2.73	54.93	2.60	2.73	2.63	2.70	2.62	2.66	2.59	2.59	2.73	2.66	2.62	2.70	2.68
2.68	2.64	2.67	2.62	2.63	54.99	2.80	2.59	2.59	2.64	2.59	2.65	2.55	2.77	2.68	2.57	2.70	2.62
2.79	2.66	2.64	2.67	2.67	2.61	54.87	2.68	2.70	2.58	2.67	2.67	2.63	2.69	2.66	2.59	2.63	2.60
2.65	2.54	2.76	2.65	2.75	2.76	2.47	54.97	2.69	2.63	2.72	2.65	2.54	2.61	2.72	2.61	2.65	2.62
2.70	2.67	2.60	2.64	2.84	2.67	2.65	2.66	54.98	2.59	2.52	2.58	2.73	2.64	2.67	2.73	2.56	2.57
2.73	2.58	2.65	2.70	2.64	2.63	2.74	2.62	2.67	54.76	2.69	2.56	2.69	2.62	2.74	2.67	2.58	2.73
2.70	2.58	2.71	2.64	2.58	2.65	2.68	2.62	2.51	2.78	55.04	2.71	2.61	2.57	2.62	2.68	2.66	2.65
2.63	2.65	2.58	2.68	2.62	2.57	2.61	2.62	2.57	2.66	2.75	55.23	2.62	2.56	2.66	2.69	2.74	2.57
2.64	2.66	2.64	2.57	2.57	2.75	2.69	2.66	2.64	2.67	2.60	2.57	55.08	2.53	2.64	2.76	2.64	2.69
2.58	2.59	2.63	2.60	2.64	2.69	2.65	2.61	2.65	2.63	2.61	2.58	2.68	54.96	2.60	2.73	2.79	2.81
2.61	2.66	2.67	2.66	2.60	2.59	2.67	2.65	2.75	2.67	2.75	2.66	2.60	2.60	54.82	2.71	2.68	2.63
2.74	2.67	2.53	2.65	2.65	2.61	2.66	2.73	2.72	2.61	2.62	2.72	2.62	2.76	2.70	54.82	2.51	2.65
2.67	2.65	2.61	2.58	2.70	2.66	2.68	2.71	2.68	2.61	2.65	2.60	2.70	2.59	2.74	2.70	54.87	2.58
2.59	2.64	2.66	2.63	2.52	2.64	2.67	2.69	2.65	2.64	2.66	2.76	2.62	2.72	2.61	2.70	2.71	54.88

Σ_u

	BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC	CS10
BAC	31.92	3.85	8.56	-0.05	7.85	3.71	5.99	6.69	5.76	0
BSC	3.85	25.47	3.85	0.01	4.02	2.34	3.14	3.63	2.81	0
Citi	8.56	3.85	32.76	-0.04	7.49	3.80	6.21	6.67	6.23	0
GS	-0.05	0.01	-0.04	14.45	-0.03	0.01	-0.06	-0.02	-0.03	0
JPM	7.85	4.02	7.49	-0.03	35.26	4.10	5.10	7.12	5.23	0
MER	3.71	2.34	3.80	0.01	4.10	25.42	2.96	3.63	3.08	0
MS	5.99	3.14	6.21	-0.06	5.10	2.96	27.16	4.76	4.32	0
WB	6.69	3.63	6.67	-0.02	7.12	3.63	4.76	29.20	4.88	0
WFC	5.76	2.81	6.23	-0.03	5.23	3.08	4.32	4.88	27.54	0
CS10	0	0	0	0	0	0	0	0	0	72.93

Table 3: Reference model parameters estimated using the MCMC procedure of Appendix A. The transition probability matrices Ω_f and Ω_c as well as the covariance matrix Σ_u are reported in percentage terms. The parameters are estimated using 10000 draws from the Gibbs sampler, with a 1000 draw burn-in period.

	Pre-crisis	Pre- Bear Stearns	Pre- Lehman Brothers
θ^{-1}	0.52	0.43	0.51
St.Dev.	0.026	0.031	0.029

Table 4: Estimates of the misspecification parameter θ^{-1} and the standard deviation of the parameter using different periods of observations of CDS rates. Draws are made using the Metropolis-Hastings procedure.

	BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC
July 31 2007	36.2	161.7	37.2	81.2	55	74.7	75.2	39.5	35.9
	29.6	51.3	34.9	71.2	54.7	73.8	73.2	39.3	38.7
August 31 2007	39.7	135.7	45.5	68.8	45.4	71.7	68.8	39.4	35
	22.1	128.7	41.4	36.5	43.2	72.2	64.7	32.3	32.2
March 31 2008	86.8	122.7	138.2	115	87.5	195.8	153.9	142.8	80.8
	65.1	157.0	128.62	133.9	85.3	109.7	93.4	83.2	90.9
September 30 2008	170	143.3	301.7	452.5	143.8	410.8	1022	385.8	170
	102.7	144.5	486.4	482.8	152.6	380.03	317.3	386.5	138.1
October 31 2008	133.1	120.2	197.6	313.3	119.9	216.2	413.3	121	97.2
	133.3	161.3	178.8	380.7	146.6	203.5	416.26	131.3	93.1

Table 5: Observed 5 year CDS rates and model-implied 5 year CDS rates at four different dates. The reference model parameters are estimated using the Gibbs sampling procedure of Appendix A and the misspecification parameter θ^{-1} using the Metropolis-Hastings procedure

Panel A: Misspecified Model

	BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC
July 31 2007	9.47	8.66	9.89	9.42	7.43	10.73	11.28	11.07	10.93
	28.16	-24.32	-22.76	134.71	-10.12	-14.55	-1.26	-14.59	-14.61
August 31 2007	7.39	11.48	12.88	3.95	8.31	12.59	11.46	13.00	12.84
	-21.92	32.53	30.19	-58.01	11.95	17.41	1.63	17.44	17.50
March 31 2008	7.55	0.00	10.32	3.84	5.62	10.13	9.58	10.60	10.42
	2.06	-100.00	-19.46	-4.24	-31.99	-19.24	-16.12	-18.17	-18.51
September 30 2008	9.08	–	9.59	8.79	7.07	10.43	10.99	10.77	10.63
	20.42	–	-26.21	120.86	-14.97	-15.87	-2.80	-15.87	-15.91
October 31 2008	7.47	–	12.98	3.67	8.06	12.21	11.15	12.63	12.47
	-17.77	–	35.38	-58.28	13.95	17.08	1.39	17.27	17.27

Panel B: Reference Model

	BAC	BSC	Citi	GS	JPM	MER	MS	WB	WFC
July 31 2007	16.47	16.47	16.47	16.47	16.47	16.47	16.47	16.47	16.47
	0.04	0.04	0.04	0.04	0.04	0.04	0.04	0.04	0.04
August 31 2007	16.46	16.46	16.46	16.46	16.46	16.46	16.46	16.46	16.46
	-0.03	-0.03	-0.03	-0.03	-0.03	-0.03	-0.03	-0.03	-0.03
March 31 2008	11.23	0.00	11.26	11.23	11.21	11.19	11.21	11.22	11.21
	-31.74	-100.00	-31.63	-31.79	-31.88	-31.98	-31.90	-31.81	-31.87
September 30 2008	11.23	–	11.25	11.25	11.26	11.22	11.21	11.22	11.22
	-28.05	–	-27.97	-27.99	-27.92	-28.12	-28.18	-28.13	-28.15
October 31 2008	11.32	–	11.35	11.34	11.35	11.32	11.31	11.32	11.31
	0.81	–	0.85	0.84	0.79	0.84	0.85	0.87	0.85

Table 6: Expected time to default and the percentage change in the expected time to default relative to previous month for different financial institutions. Panel A: expected time to default perceived by the misspecification-averse agent. Panel B: expected time to default under the reference model. The reference model parameters are estimated using the Gibbs sampling procedure of Appendix A and the misspecification parameter θ^{-1} using the Metropolis-Hastings procedure.

July 31 2007

	Data	0	0.5	1	2	3	4	5
BAC	36.2	1.3	1.4	1.6	1.7	1.7	1.5	1.2
BSC	161.7	1.6	1.6	1.6	1.6	1.4	1.0	0.7
Citi	37.2	1.4	1.5	1.5	1.6	1.5	1.2	0.9
GS	81.2	1.4	1.4	1.3	1.2	1.0	0.8	0.5
JPM	55	1.2	1.8	2.7	5.7	10.8	17.9	25.9
MER	74.7	1.4	1.5	1.6	1.7	1.6	1.4	1.0
MS	75.2	1.4	1.5	1.6	1.8	1.7	1.5	1.1
WB	39.5	1.4	1.5	1.6	1.7	1.7	1.5	1.1
WFC	35.9	1.4	1.5	1.6	1.7	1.7	1.4	1.1

August 31 2007

	Data	0	0.5	1	2	3	4	5
BAC	39.7	1.9	2.1	2.2	2.5	2.5	2.2	1.7
BSC	135.7	2.1	2.1	2.1	2.0	1.7	1.3	0.9
Citi	45.5	2.0	2.1	2.2	2.2	2.1	1.7	1.2
GS	68.8	2.0	2.0	1.9	1.8	1.5	1.1	0.7
JPM	45.4	1.8	2.6	3.9	8.2	15.6	25.9	37.5
MER	71.7	2.0	2.1	2.2	2.4	2.3	2.0	1.5
MS	68.8	2.0	2.1	2.3	2.5	2.5	2.2	1.6
WB	39.4	2.0	2.1	2.3	2.5	2.5	2.1	1.6
WFC	35	2.0	2.1	2.3	2.4	2.4	2.0	1.5

March 31 2008

	Data	0	0.5	1	2	3	4	5
BAC	86.8	19.0	19.4	19.6	19.2	17.3	13.9	9.9
BSC	122.7	861.4	795.8	729.6	591.7	445.4	301.7	181.3
Citi	138.2	19.5	19.3	18.9	17.4	14.7	11.1	7.4
GS	115	20.2	18.8	17.4	14.2	10.7	7.2	4.2
JPM	87.5	18.6	26.2	36.5	67.9	116.0	176.8	237.7
MER	195.8	20.4	20.6	20.6	19.6	17.2	13.4	9.3
MS	153.9	19.9	20.2	20.4	19.8	17.7	14.1	9.9
WB	142.8	19.8	20.1	20.2	19.5	17.3	13.7	9.5
WFC	80.8	19.7	19.9	20.0	19.1	16.8	13.2	9.2

September 30 2008

	Data	0	0.5	1	2	3	4	5
BAC	170	19.2	19.6	19.8	19.5	17.5	14.1	10.0
BSC	143.3	19.8	18.8	17.7	14.9	11.6	8.1	4.9
Citi	301.7	19.8	19.6	19.2	17.7	15.0	11.3	7.5
GS	452.5	19.7	18.4	17.0	13.9	10.5	7.0	4.1
JPM	143.8	17.5	24.6	34.3	63.9	109.2	166.4	223.8
MER	410.8	19.7	19.9	19.8	18.9	16.6	13.0	9.0
MS	1022	20.0	20.3	20.5	19.9	17.8	14.2	9.9
WB	385.8	20.1	20.4	20.5	19.7	17.5	13.9	9.7
WFC	170	19.7	20.0	20.0	19.2	16.9	13.3	9.2

Table 7: Observed five year CDS and implied CDS for different levels of the risk aversion coefficient, γ, at four different dates. The reference model parameters are estimated using the Gibbs sampling procedure of Appendix A.

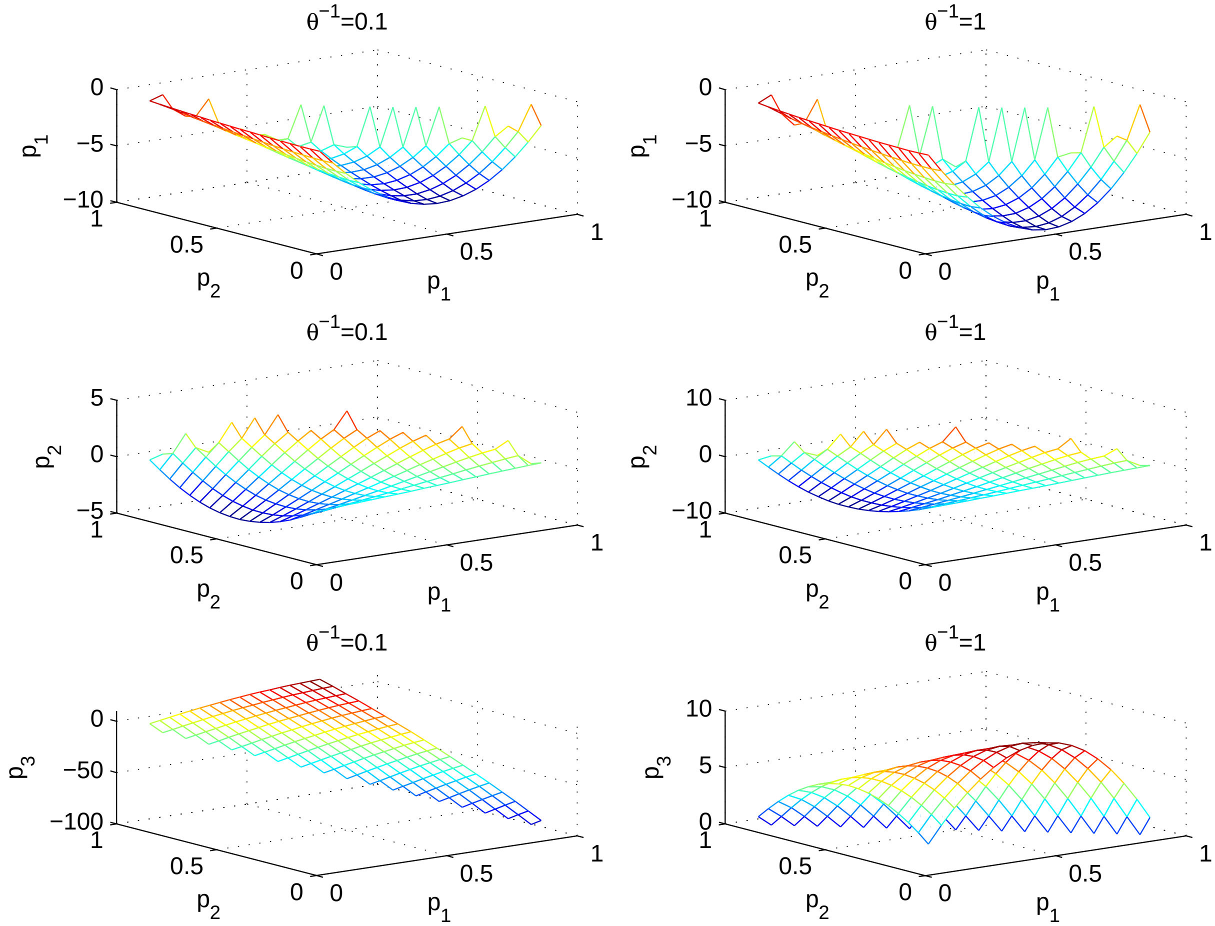

Figure 3: Differences between the reference conditional probabilities and the misspecification-implied conditional probabilities. Left column: $\theta = 0.1$; right column: $\theta = 1$. The parameters used are those of Section 3.3.1.

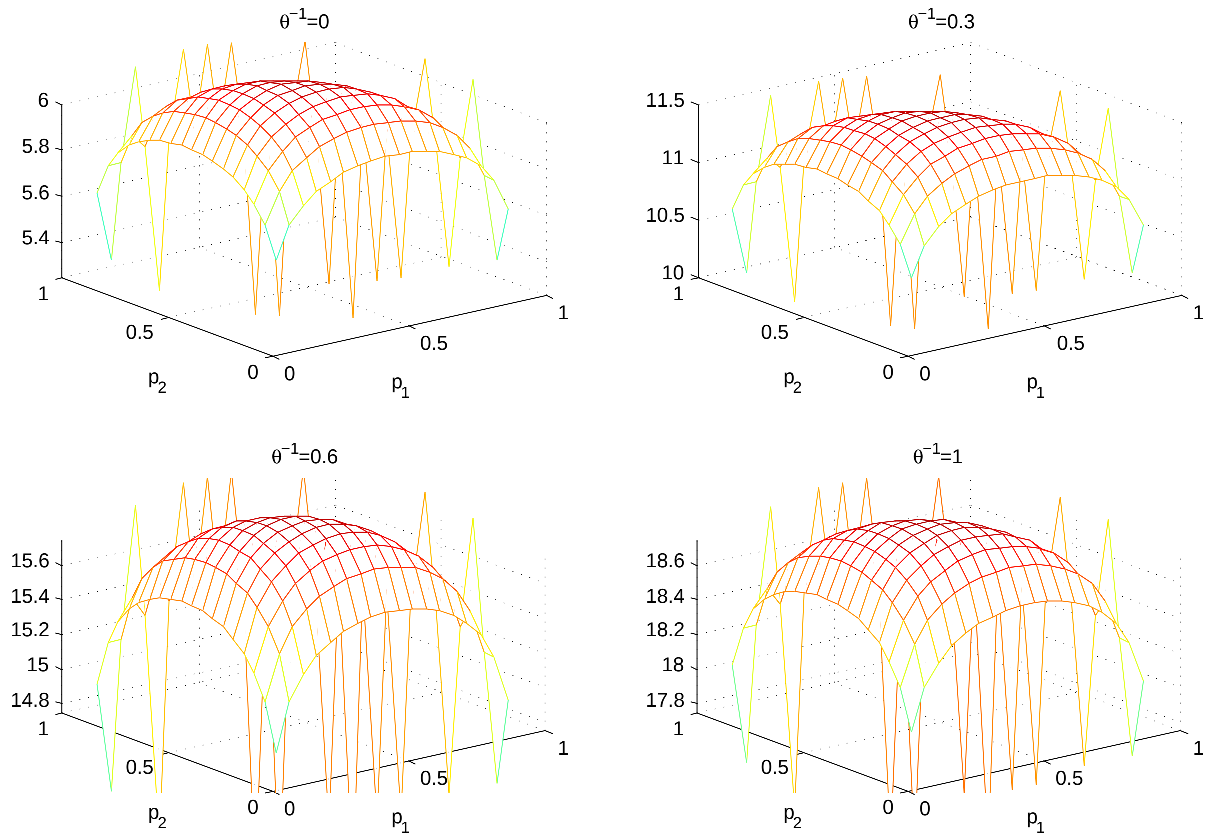

Figure 4: 5 year CDS rate for various values of the misspecification parameter θ^{-1} when no defaults have occurred. The parameters used are those of Section 3.3.1.

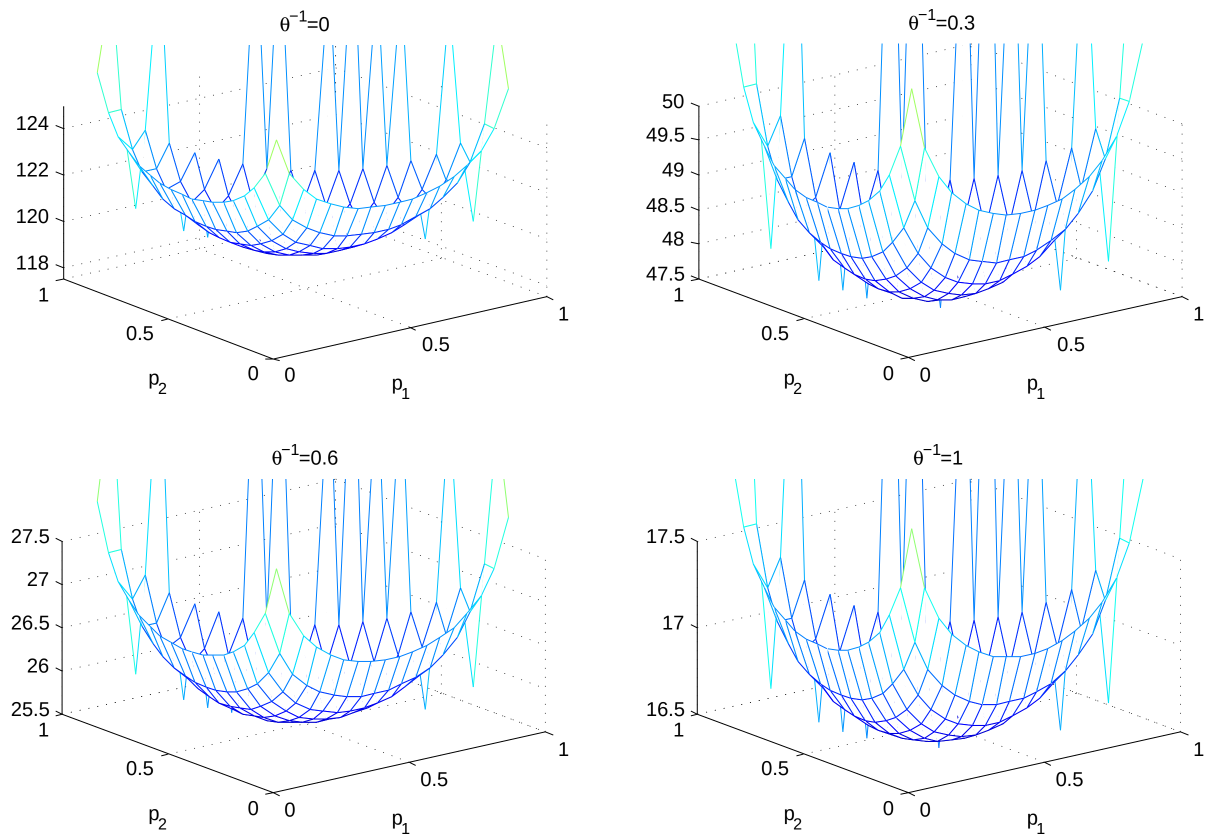

Figure 5: Percentage change in the 5 year CDS rate for various values of the misspecification parameter θ^{-1} after the other firm defaults. The parameters used are those of Section 3.3.1.

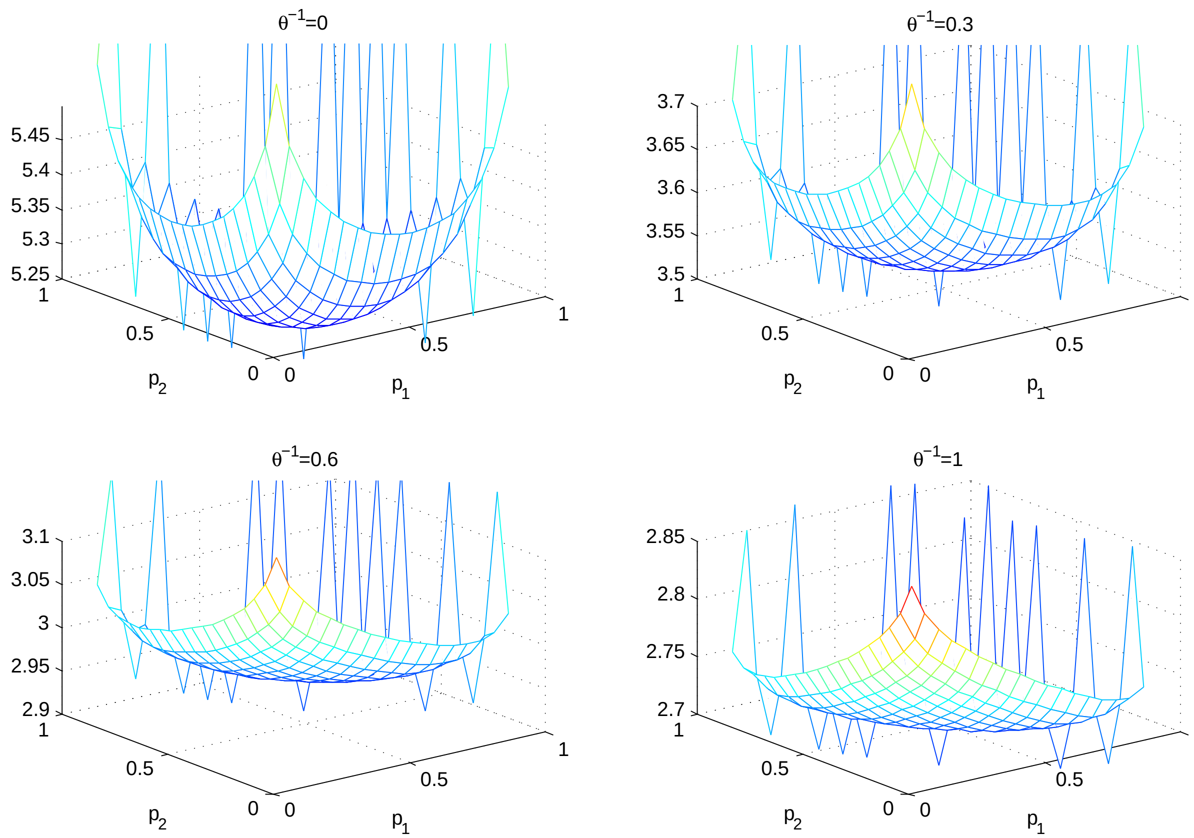

Figure 6: Expected time to default for various values of the misspecification parameter θ^{-1} when no defaults have occurred. The parameters used are those of Section 3.3.1.

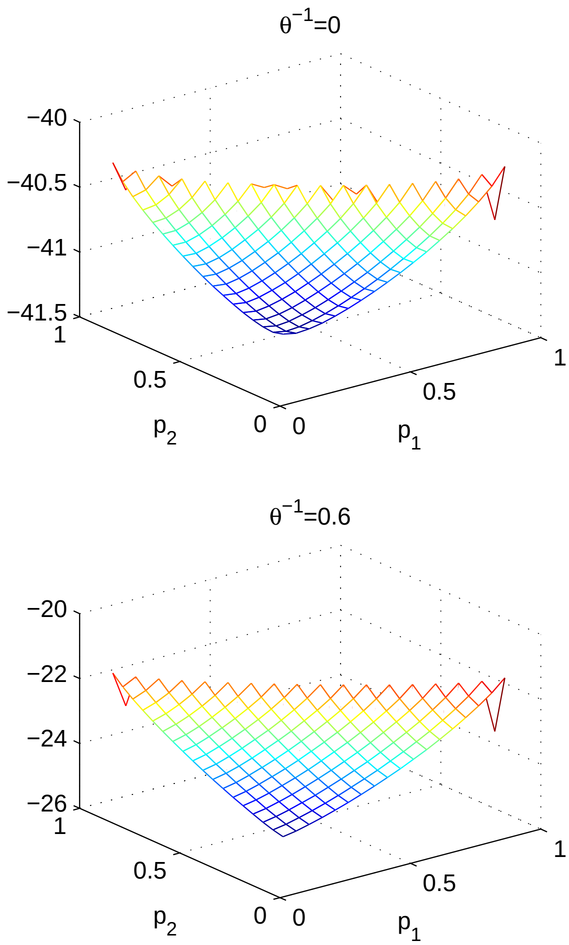

Figure 7: Percentage change in the expected time to default for various values of the misspecification parameter θ^{-1} after the other firm defaults. The parameters used are those of Section 3.3.1.

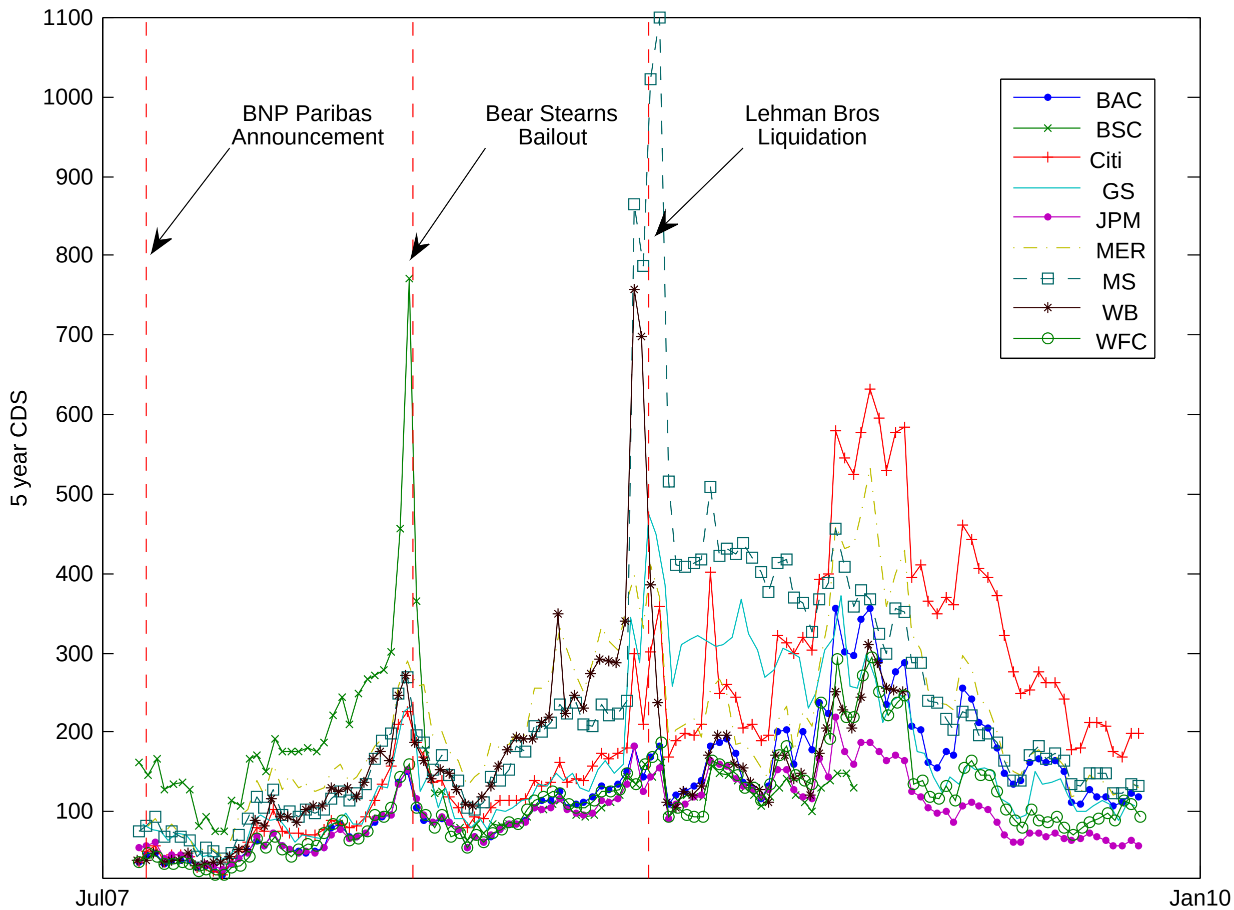

Figure 8: Evolution of the five year CDS spreads for financial institutions over the course of the crisis. Data source: Datastream

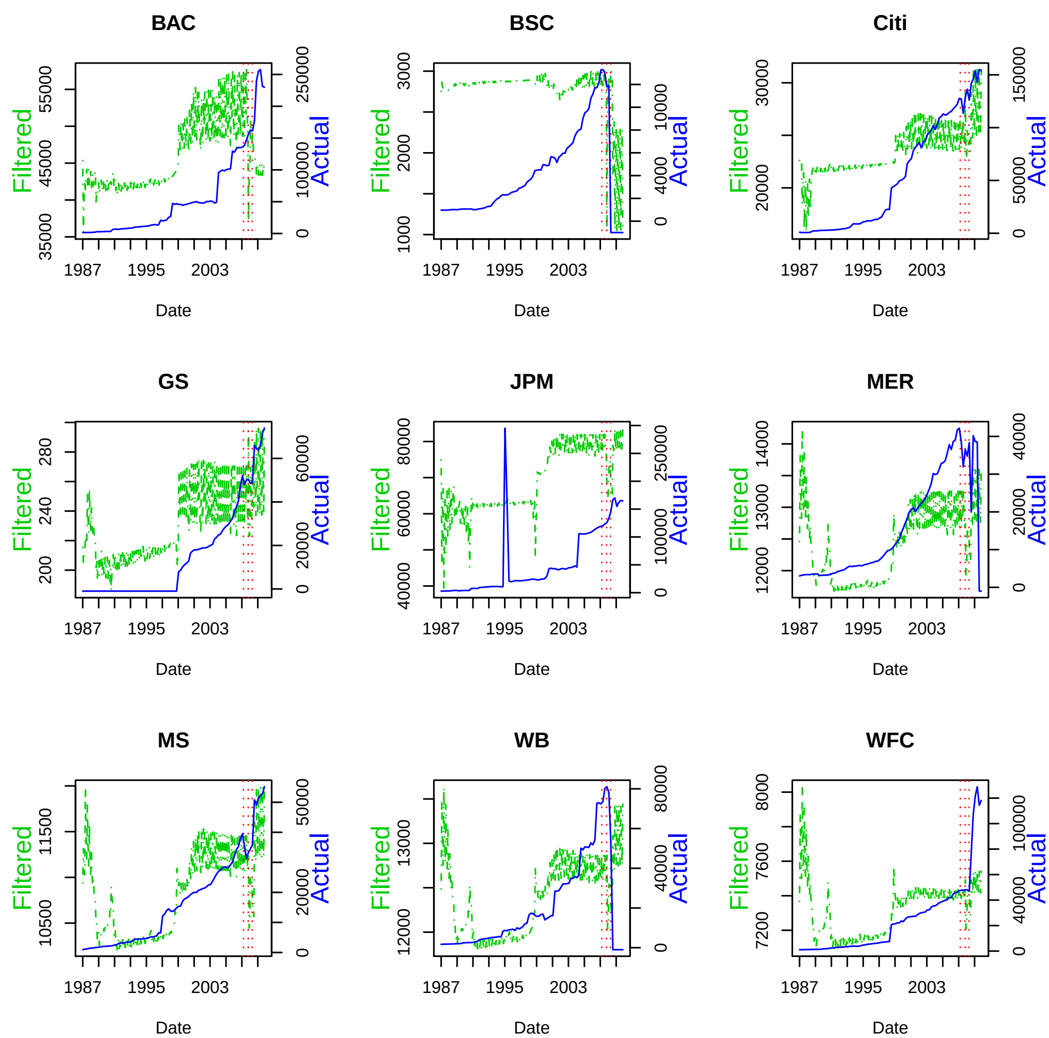

Figure 9: Filtered estimates of fundamental asset value evolution (left-hand scale) and the observed book value evolution (right-hand scale) for different financial institutions. Parameters are estimated using the Gibbs Sampling procedure of Appendix A with 10000 draws and 1000 draw burn-in. The different states are weighted using the reference model conditional probabilities.

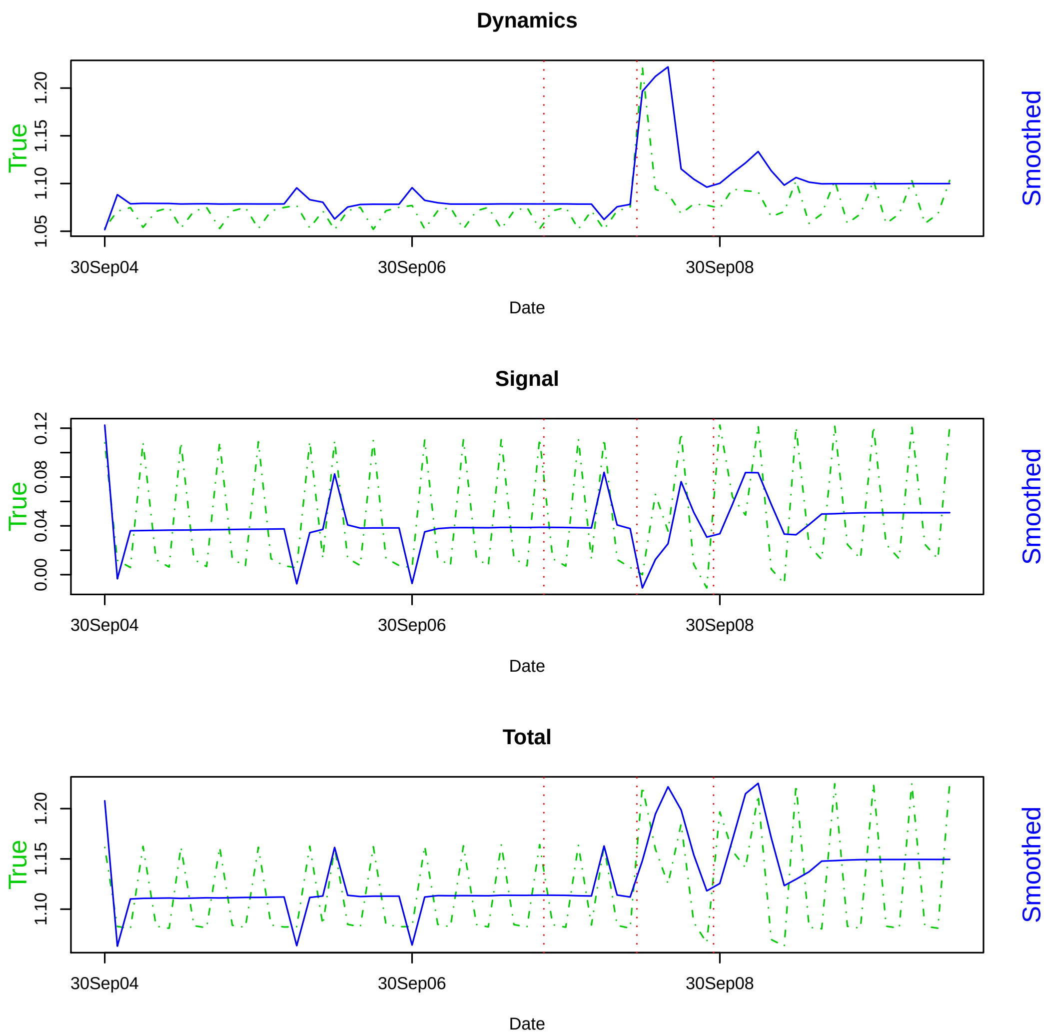

Figure 10: Components of relative entropy between the reference and misspecified models over time. Upper panel: entropy due to misspecification of the joint signals and states dynamics; central panel: entropy due to misspecification of the current period conditional probability; lower panel: total entropy. The right hand scale in each panel is the three month moving average of the corresponding entropy measure. The misspecification parameter, θ^{-1} is estimated using a Metropolis-Hastings algorithm and observations of CDS spreads.

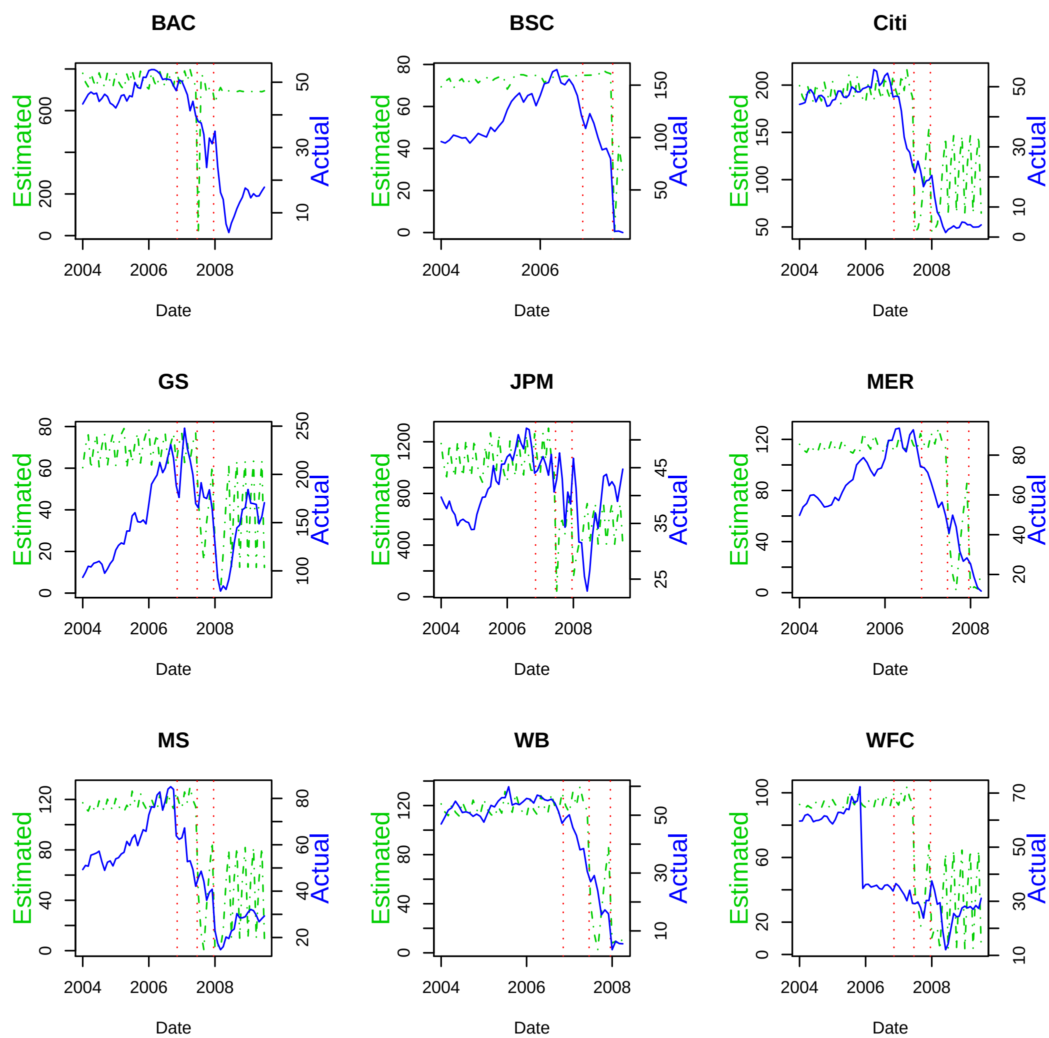

Figure 11: Value of equity under the misspecified model (left-hand scales) and the observed value of equity (right-hand scale) of the different financial institutions over time. The misspecification parameter, θ^{-1} is estimated using a Metropolis-Hastings algorithm and observations of CDS spreads.