Lucas, André; Schwaab, Bernd; Zhang, Xin

Working Paper
Modeling financial sector joint tail risk in the euro area

Sveriges Riksbank Working Paper Series, No. 308

Provided in Cooperation with:
Central Bank of Sweden, Stockholm

Suggested Citation: Lucas, André; Schwaab, Bernd; Zhang, Xin (2015) : Modeling financial sector joint tail risk in the euro area, Sveriges Riksbank Working Paper Series, No. 308, Sveriges Riksbank, Stockholm

This Version is available at:
http://hdl.handle.net/10419/129723

Terms of use:
 Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.
Modeling financial sector joint tail risk in the euro area

André Lucas, Bernd Schwaab and Xin Zhang

June 2015
WORKING PAPERS ARE OBTAINABLE FROM

Sveriges Riksbank • Information Riksbank • SE-103 37 Stockholm
Fax international: +46 8 787 05 26
Telephone international: +46 8 787 01 00
E-mail: info@riksbank.se

The Working Paper series presents reports on matters in the sphere of activities of the Riksbank that are considered to be of interest to a wider public. The papers are to be regarded as reports on ongoing studies and the authors will be pleased to receive comments.

The views expressed in Working Papers are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Executive Board of Sveriges Riksbank.
Modeling financial sector joint tail risk in the euro area

André Lucas†, Bernd Schwaab‡, Xin Zhang§

Sveriges Riksbank Working Paper Series No. 308

June 2015

Abstract
We develop a novel high-dimensional non-Gaussian modeling framework to infer measures of conditional and joint default risk for many financial sector firms. The model is based on a dynamic Generalized Hyperbolic Skewed-$t$ block-equicorrelation copula with time-varying volatility and dependence parameters that naturally accommodates asymmetries, heavy tails, as well as non-linear and time-varying default dependence. We apply a conditional law of large numbers in this setting to define joint and conditional risk measures that can be evaluated quickly and reliably. We apply the modeling framework to assess the joint risk from multiple defaults in the euro area during the 2008–2012 financial and sovereign debt crisis. We document unprecedented tail risks during 2011–2012, as well as their steep decline after subsequent policy actions.

Keywords: dynamic equicorrelation; generalized hyperbolic distribution; law of large numbers; large portfolio approximation.

JEL classifications: C32, G21.
1 Introduction

In this paper we develop a novel high-dimensional non-Gaussian modeling framework to infer conditional and joint risk measures for many financial sector firms. The model is based on a dynamic Generalized Hyperbolic Skewed-$t$ copula with time-varying volatility and dependence parameters. Such a framework naturally accommodates asymmetries, heavy tails, as well as non-linear and time-varying default dependence. To balance the need of parsimony as well as flexibility in a high-dimensional cross-section, we endow the dynamic model with a score-driven block-equicorrelation structure. We further demonstrate that a conditional law of large numbers applies in our setting, which allows us to define risk measures that can be evaluated semi-analytically and within seconds. The modeling framework allows us to assess the joint risk from multiple financial firm defaults in the euro area during the financial and sovereign debt crisis. We document unprecedented tail risks during 2011-12, as well as a sharp decline in joint (but not conditional) tail risk probabilities after a sequence of announcements by the European Central Bank (ECB) that introduced its Outright Monetary Transactions (OMT) program.\(^1\)

Since the onset of the financial crisis in 2007, financial stability monitoring has become a key priority for many central banks, in addition to their respective monetary policy mandates; see, for example, Acharya et al. (2012), and Adrian et al. (2013). The new prudential responsibilities involve the analysis of financial risks to and from a large system of interconnected financial intermediaries. The cross-sectional dimensions of these systems are typically high, even if attention is restricted to only large and systemically important institutions. Our modeling framework is directly relevant for such monitoring tasks. In addition, our framework is interesting for financial institutions and clearing houses that are required to actively set risk limits (often in real time) and maintain economic capital buffers to withstand bad risk outcomes due to exposures to a large number of credit risky counterparties. Finally, with the benefit of hindsight, an evaluation of the time variation in conditional and joint

\(^1\)See ECB (2012) and Coeuré (2013). The OMT is a non-standard monetary policy measure within which the ECB could, under certain conditions, make purchases in secondary markets of bonds issued by euro area member states.
risks allows us to assess the impact of non-standard policy measures undertaken by central banks (or other actors) on the risk of a simultaneous and widespread failure of financial intermediaries.

Our starting point for modeling time-varying joint and conditional risks is a dynamic copula framework, as also considered by, for example, Segoviano and Goodhart (2009), Christoffersen, Errunza, Langlois, and Jacobs (2012), Oh and Patton (2014), and Lucas, Schwaab, and Zhang (2014). In each case, a collection of firms is seen as a portfolio of firms whose multivariate dependence structure is inferred from equity returns or CDS data. Our current framework extends Lucas et al. (2014) in two ways. First, by considering grouped equicorrelation structures, our current framework allows us to fit a cross-sectional dimension much larger than, say, 15 firms, while retaining the ability to capture the salient data features such as skewness, fat tails, and time-varying correlations. We thus fix the drawback that estimation brakes down for many firms due to a well-known curse of dimensionality for this class of models; see Engle and Kelly (2012) for a discussion. Second, we show how to evaluate joint and conditional risk measures within seconds in the current framework. The estimation of time-varying parameters, as well as portfolio risk measures, is computationally relatively inexpensive because explicit expressions are available for the likelihood and the joint and conditional portfolio credit risk measures. Simulation-based methods are not required, but are used in our empirical study to provide points of comparison from a robustness perspective.

We apply our general framework by studying joint and conditional default probabilities for financial sector firms in the euro area, based on weekly data from January 1999 to September 2013. We consider two applications. First, using a limited sample of $N = 10$ firms, we verify that our dynamic correlations based on the block-equicorrelation assumption closely track the average correlations from a full-correlation-matrix model analysis. We demonstrate that the loss of cross-sectional dispersion in correlations matters little when evaluating joint credit risk measures, at least for our sample of firms. Finally, we verify that our semi-analytical approximations to compute joint and conditional risk measures already work well even if the cross-sectional dimension is as low as 10 firms.

We then turn to our final high-dimensional study of $N = 73$ firms. We document
unprecedented joint tail risks for a set of large euro area financial sector firms during the financial and sovereign debt crises, and also establish a clear peak of financial sector joint default risk in the summer of 2012. Based on time variation in our joint risk measures we argue that three events — a speech by the ECB president in London on 26 July 2012, the announcement of the ECB’s OMT program on 2 August 2012, and the disclosure of the OMT details on 6 September 2012 — collectively ended the most acute phase of extreme financial sector tail risks in the euro area. This is a surprising finding, as the ECB’s OMT program provides conditional and partial insurance to governments, not financial sector firms. Conditional tail probabilities did not decline as much, indicating that risk spillovers may have remained a concern. Based on the OMT’s strong impact on financial sector joint risks we conclude that the design and implementation of unconventional monetary policies and financial stability (tail) outcomes are strongly related. This finding suggests substantial scope for the coordination of monetary, macro-prudential, and bank supervision policies. This is relevant as both monetary policy as well as banking supervision will be carried out jointly by the ECB as of November 2014.

Our study relates to several directions of current research. First, we draw from a growing literature on non-Gaussian dependence modeling as well as a parallel literature on credit risk measurement and portfolio loss asymptotics. Time-varying parameter models for volatility and dependence have been considered, inter alia, by Engle (2002), Creal, Koopman, and Lucas (2011), Christoffersen et al. (2012), Engle and Kelly (2012), and Creal and Tsay (2015). At the same time, credit risk models and portfolio tail risk measures have been studied, for example, by Vasicek (1987), Lucas, Klaassen, Spreij, and Straetmans (2001, 2003), Gordy (2000, 2003), Koopman et al. (2011, 2012), and Giesecke, Spiliopoulos, Sowers, and Sirignano (2014). Combining results from both strands of literature allows us to obtain joint credit risk measures at a relatively high frequency, such as daily or weekly. Second, to balance the need for parsimony and flexibility, we consider a variant of the block equicorrelation structure for the covariance matrix in Engle and Kelly (2012); see also Christoffersen, Errunza, Jacobs, and Jin (2014) and Creal and Tsay (2015) for applications with DECO and grouped dependence structures, respectively. The version of the block equicorrelation model developed in this
paper still allows us to draw upon the machinery developed in the credit portfolio tail risk literature mentioned earlier. Third, an additional strand of literature investigates joint and conditional default dependence from a financial stability perspective, see, for example, Hartmann et al. (2007), Acharya et al. (2012), and Suh (2012). Finally, to introduce time-variation into our econometric model specification, we endow our model with observation-driven dynamics based on the score of the conditional predictive log-density. Score-driven time-varying parameter models are an active area of research, see for example Creal et al. (2011), Creal et al. (2013), Harvey (2013), Oh and Patton (2014), Creal et al. (2014), and more.²

The two papers that are most closely related to ours are Oh and Patton (2014), and Christoffersen, Jacobs, Jin, and Langlois (2014). Oh and Patton (2014) propose a class of dynamic copula factor models for high dimensions, which facilitates the estimation of a wide variety of systemic risk measures. Firms in their framework load on a common factor that has the skewed Student’s $t$ density of Hansen (1994). Additional idiosyncratic shocks are modeled by a symmetric-$t$ distribution. They use their model for studying the risk from a large number of U.S. non-financial corporates. Time-varying dependence is modeled in a score-driven way, as in this paper. Our studies differ in that we use the GHST distribution for both the marginal and the copula modeling; this means that our factor copula has a non-linear two factor rather than a linear one-factor structure. Moreover, our application focusses on the euro area financial sector during the financial and sovereign debt crisis, and on the impact evaluation of central bank unconventional monetary policy measures during this time. Finally, we introduce and discuss the efficient evaluation of semi-analytic risk measures. Christoffersen et al. (2014) study diversification benefits among U.S. corporates, and use Hansen’s (1994) distribution for the innovations in univariate GARCH models. Their GHST copula is the same as used in our paper. Our paper differs in that we provide a score-driven approach to the modeling of dynamic dependence, which is particularly attractive in a non-Gaussian context when data is fat tailed and skewed; see Zhang, Creal, Koopman,

²We refer to http://www.gasmodel.com for an extensive enumeration of recent work in this area. Computer code for this paper will be available from this source.
and Lucas (2011) and Blasques, Koopman, and Lucas (2015). We do not rely exclusively on composite likelihood methods, but employ them only to provide alternative points of comparison from a robustness perspective. Composite likelihood techniques are feasible in high dimensions, but also statistically inefficient. Rather, we proceed by proposing block equicorrelation models, at least for settings where prior information is available on which subsets of firms are likely to move together as a group, and risks need to be evaluated in a computationally efficient way. We explicitly highlight what information is lost when moving from a full to an equicorrelation copula structure. Finally, as for Oh and Patton (2014), we introduce semi-analytic risk measures and focus our application on the euro area sovereign debt crisis and non-standard monetary policy measures.

The remainder of the paper is set up as follows. Section 2 introduces our statistical framework, the dynamic Generalized Hyperbolic Skewed-$t$ block equicorrelation model, and discusses parameter estimation. Section 3 demonstrates how a conditional law of large numbers can be applied to reliably and quickly compute portfolio risk measures in a GHST factor copula setting. Section 4 applies the modeling framework to the euro area financial sector during the financial and sovereign debt crisis. Section 5 concludes. A supplementary Web Appendix presents proofs and additional results.

2 Statistical model

2.1 The dynamic Generalized Hyperbolic Skewed $t$ copula model

Following Lucas et al. (2014), we consider a copula model based on the Generalized Hyperbolic Skewed $t$ (GHST) distribution. Let

\[ y_{it} = (\varsigma_t - \mu_i)\gamma_i + \sqrt{\varsigma_t} \tilde{\Sigma}_{it}^{1/2} \epsilon_t, \quad i = 1, \ldots, N, \]

where $y_t = (y_{1t}, \ldots, y_{Nt})'$ is a vector of firm-specific log asset values, $\varsigma_t \in \mathbb{R}^+$ is an inverse-Gamma distributed common risk factor that affects all firms simultaneously, $\varsigma_t \sim IG(\frac{\alpha}{2}, \frac{\beta}{2})$, $\gamma \in \mathbb{R}^N$ is a vector controlling the skewness of the copula, $\tilde{\Sigma}_{it}$ is the $i$th row of the GHST
copula scale matrix $\tilde{\Sigma}_t \in \mathbb{R}^{N \times N}$, and $\epsilon_t \in \mathbb{R}^N$ is a vector of standard normally distributed risk factors. We assume that the two random variables $\varsigma_t$ and $\epsilon_t$ are independent and set $\mu_t = \mathbb{E}[\varsigma_t] = \nu/(\nu - 2)$, such that $y_{it}$ has zero mean if $\nu > 2$. Matrix $\tilde{\Sigma}_t^{1/2}$ is the Choleski matrix square root of the scale matrix $\tilde{\Sigma}_t$.

In our copula framework, a firm defaults if its (unobserved) log asset value $y_{it}$ falls below its default threshold $y_{it}^\ast$. The cross sectional dependence in defaults captured by equation (1) thus stems from two sources: common exposures to the normally distributed risk factors $\epsilon_t$ as captured by the time-varying matrix $\tilde{\Sigma}_t$; and an additional common exposure to the scalar risk factor $\varsigma_t$. The former captures connectedness through correlations, while the latter captures such effects through the tail-dependence of the copula. To see this, note that if $\varsigma_t$ is non-random, the first term in (1) drops out of the equation and there is zero tail dependence. Conversely, if $\varsigma_t$ is large, all asset values are affected at the same time, making joint defaults of two or more firms more likely.

Earlier applications of the GHST distribution to financial and economic data include, for example, Mencía and Sentana (2005), Aas and Haff (2006), and Oh and Patton (2014). Alternative skewed $t$ distributions have been proposed as well, such as Branco and Dey (2001), Gupta (2003), Azzalini and Capitanio (2003), and Bauwens and Laurent (2005); see also the overview of Aas and Haff (2006). The GHST distribution closely connects to a continuous-time finance literature utilizing Lévy processes for stock price processes and firm asset values; see Bibby and Sørensen (2003) for a survey.

We denote the one-year-ahead default probability for firm $i$ at time $t$ as $p_{it}$, such that

$$p_{it} = \Pr[y_{it} < y_{it}^\ast] = F_{it}(y_{it}^\ast) \iff y_{it}^\ast = F_{it}^{-1}(p_{it}),$$

where $F_{it}$ is the univariate GHST cumulative distribution function (cdf) of $y_{it}$. In our application, we assume that we observe $p_{it}$ as the expected default frequency of firm $i$ reported at time $t$ by Moody’s Analytics (formerly Moody’s KMV). Instead of focusing on the individual default probabilities $p_{it}$, however, our focus is on the time-varying joint probabilities, such as $\Pr[y_{it} < y_{it}^\ast, y_{jt} < y_{jt}^\ast]$, and on conditional probabilities such as $\Pr[y_{it} < y_{it}^\ast | y_{jt} < y_{jt}^\ast]$, for
firms $i \neq j$. Below, we first develop a dynamic version of the GHST copula model. We then consider a dynamic (block)-equicorrelation version of the model in the spirit of Engle and Kelly (2012), which turns out to be particularly useful to study joint and conditional default probabilities in a parsimonious way for large dimensional systems.

2.2 The dynamic GHST model

To describe the dynamics of the scale parameter $\tilde{\Sigma}_t$ in the GHST model (1), we use the generalized autoregressive score (GAS) dynamics as proposed in Creal et al. (2011, 2013); see also Harvey (2013). These dynamics easily adapt to the skewed and fat-tailed nature of the GHST density and improve the stability of dynamic volatility and correlation estimates; see Blasques et al. (2015). Our version of the model is different from that in Lucas et al. (2014) due to a different parameterization. We consider the scale matrix (rather than the covariance matrix) in order to fully employ the block equicorrelation structure later on for our conditional law of large numbers result.

To derive the score dynamics for the GHST model, we need the conditional density of $y_t$, which we parameterize as

$$p(y_t; \tilde{\Sigma}_t, \gamma, \nu) = \frac{2(\nu/2)^{\nu/2}}{\Gamma(\nu/2)} \frac{K_{(\nu+N)/2} \left( \sqrt{d(y_t) \cdot d(\gamma)} \right)}{(d(y_t)/d(\gamma))^{(\nu+N)/4}} e^{\gamma\tilde{\Sigma}_t^{-1}(y_t-\tilde{\mu})},$$

$$d(y_t) = \nu + (y_t - \tilde{\mu})' \tilde{\Sigma}_t^{-1}(y_t - \tilde{\mu}),$$

$$d(\gamma) = \gamma' \tilde{\Sigma}_t^{-1} \gamma, \quad \tilde{\mu} = -\frac{\nu}{\nu - 2},$$

where $K_a(b)$ is the modified Bessel function of the second kind; see Bibby and Sørensen (2003). The parameters $\gamma = (\gamma_1, \ldots, \gamma_N)' \in \mathbb{R}^N$ and $\nu \in \mathbb{R}^+$ are the skewness and degrees of freedom (or kurtosis) parameter, respectively, while $\tilde{\mu}$ and $\tilde{\Sigma}_t$ denote the location vector and scale matrix, respectively. Note that if $y_t$ has a multivariate GHST distribution with parameters $\tilde{\mu}$, $\tilde{\Sigma}$, $\gamma$, and $\nu$ as given in (3), then $Ay_t + b$ for some matrix $A$ and vector $b$ also has a GHST distribution, with parameters $A\tilde{\mu} + b$, $A\tilde{\Sigma}A'$, $A\gamma$, and $\nu$. In particular, the marginal distributions of $y_{it}$ also have a GHST distribution. The GHST density (3) nests the
symmetric-$t$ ($\gamma = 0$) and multivariate normal ($\gamma = 0$ and $\nu \to \infty$) distributions as special cases.

We parameterize the time-varying matrix $\hat{\Sigma}_t$ as in Engle (2002), i.e.,

$$\hat{\Sigma}_t = D(f_t)\hat{R}(f_t)D(f_t),$$

where $f_t$ is a vector of time varying parameters, $D(f_t)$ is a diagonal matrix holding the scale parameters of $y_{it}$, and $\hat{R}(f_t)$ captures the dependence parameters. In our current copula set-up, we use univariate models for $D(f_t)$, and the multivariate model for $\hat{R}(f_t)$. The Web Appendix provides further detail on our univariate volatility modeling approach for $D(f_t)$.

For the remainder of this section, we concentrate on the matrix $\hat{R}(f_t)$.

Following Creal et al. (2011, 2013), we endow $f_t$ with score-driven dynamics using the derivative of the log conditional observation density (3). The transition dynamics for $f_t$ are given by

$$f_{t+1} = \tilde{\omega} + \sum_{i=0}^{p-1} A_i s_{t-i} + \sum_{j=0}^{q-1} B_j f_{t-j},$$

$$s_t = S_t \nabla_t, \quad \nabla_t = \partial \ln p(y_t|\mathcal{F}_{t-1}; f_t, \theta)/\partial f_t,$$

where $\tilde{\omega} = \tilde{\omega}(\theta)$ is a vector of fixed intercepts, $A_i = A_i(\theta)$ and $B_j = B_i(\theta)$ are fixed parameter matrices that depend on the vector $\theta$ containing all time invariant parameters in the model, and $S_t$ a scaling function.

The key element in (7) is the scaled score $s_t$. We can show that if $y_t$ has a zero mean GHST distribution $p(y_t; \hat{\Sigma}_t, \gamma, \nu)$ and the time-varying scale matrix is driven by (7)-(8), then the score is given by

$$\nabla_t = \Psi_t^t H_t^t \text{vec} \left( w_t \cdot (y_t - \hat{\mu}) (y_t - \hat{\mu})' - 0.5\hat{\Sigma}_t - \gamma (y_t - \hat{\mu})' - \tilde{\omega} \cdot \gamma \gamma' \right),$$

(9)
where

\[ w_t = \nu + N \frac{k'_{0.5(\nu+N)}(\sqrt{d(y_t)d(\gamma)})}{2\sqrt{d(y_t)/d(\gamma)}}, \]

\[ \bar{w}_t = \nu + N \frac{k'_{0.5(\nu+N)}(\sqrt{d(y_t)d(\gamma)})}{2\sqrt{d(\gamma)/d(y_t)}}, \]

\[ H_t = \hat{\Sigma}_t^{-1} \otimes \hat{\Sigma}_t^{-1}, \quad \Psi_t = \frac{\partial \text{vec}(\hat{\Sigma}_t)^{\prime}}{\partial f_t}, \]

and where \( k_{\nu}(\cdot) = \ln K_{\nu}(\cdot) \) with first derivative \( k'_{\nu}(\cdot) \). The matrices \( \Psi_t \) and \( H_t \) are time-varying and parameterization-specific; both matrices depend on \( f_t \) but not on the data \( y_t \). We refer to the Web Appendix for a derivation of (9).

Equation (8) reveals the key feature of the score-driven specification. In essence, the score-driven mechanism takes a Gauss-Newton improvement step for the scale matrix to better fit the most recent observation. Equation (8) shows that \( f_t \) reacts to deviations between \( \hat{\Sigma}_t \) and the observed \( (y_t - \tilde{\mu})(y_t - \tilde{\mu})' \). The reaction is asymmetric if \( \gamma \neq 0 \), in which case there is also a reaction to the level \( (y_t - \tilde{\mu}) \) itself. The reaction to \( (y_t - \tilde{\mu})(y_t - \tilde{\mu})' \) is modified by the weight \( w_t \). If \( \nu < \infty \), the GHST distribution is fat tailed and the weight decreases in the Mahalanobis distance \( d(y_t) \); compare the discussion for the symmetric Student’s t case in Creal et al. (2011). This feature gives the model a robustness flavor in that incidental large values of \( y_t \) have a limited impact on future volatilities and correlations. The remaining expressions for \( H_t \) and \( \Psi_t \) only serve to transform the dynamics of the covariance matrix in (6) into the dynamics of the unobserved factor \( f_t \).

To scale the score in (8) we set the scaling matrix \( S_t \) equal to the inverse conditional Fisher information matrix of the symmetric Student’s t distribution; see the Web Appendix for details. Zhang et al. (2011) demonstrate that this choice of scaling matrix results in a stable model that outperforms alternative models if the data are fat-tailed and skewed. We do not use the inverse conditional Fisher information matrix of the multivariate GHST distribution because it is not available in closed form.
2.3 Dynamic block-equicorrelation structure

As we want to use our model in the context of a large cross-sectional dimension to describe the joint (tail) risk dynamics in a large system of financial institutions, we refrain from modeling all dependence parameters in \( \tilde{R}(f_t) \) individually. Instead, we adopt the approach of Engle and Kelly (2012) and impose a block-equicorrelation structure on the matrix \( \tilde{R}(f_t) \).

By limiting the number of free parameters, we considerably facilitate the estimation process while retaining the ability to capture dynamic patterns in the dependence structure among financial firms, particularly in times of stress.

For the single block equicorrelation model, we assume that \( \tilde{\Sigma}_t \) takes the form

\[
\tilde{\Sigma}_t = (1 - \rho_t^2)I_N + \rho_t^2 \ell_N \ell_N',
\]

where \( \rho_t = (1 + \exp(-f_t))^{-1} \in (0, 1) \), and \( \ell_N \) is a \( N \times 1 \) vector of ones. In this case, the expressions for \( \Psi_t \) and \( H_t \) simplify considerably. In particular,

\[
\Psi_t = \frac{\partial \text{vec}(\tilde{\Sigma}_t)'}{\partial f_t} = (\ell_N' - \text{vec}(I_N)) \frac{2\exp(-f_t)}{(1 + \exp(-f_t))^3},
\]

which implies that the score \( \nabla_t \) reduces to a scalar process over time. We refer to the Web Appendix for a derivation. We can easily generate the equicorrelation structure (10) from model (1) by specifying

\[
y_{it} = (\varsigma_t - \mu_c) \gamma_i + \sqrt{\varsigma_t} \left( \rho_t \kappa_t + \sqrt{1 - \rho_t^2} u_{it} \right),
\]

where \( \kappa_t \) and \( u_{it} \) are two independent standard normal random variables. Note that (12) is a special case of (1). The logistic parameterization \( \rho_t = (1 + \exp(f_t))^{-1} \) forces the correlation parameter to be in the unit interval, irrespective of the value of \( f_t \in \mathbb{R} \). While the original equicorrelation specification of Engle and Kelly (2012) also allows for (slightly) negative equicorrelations, such values are typically unrealistic in the type of applications we consider later on. The parameterization with equicorrelation parameter \( \rho_t^2 > 0 \) therefore suffices for our current purposes.
A 2-block equicorrelation model may be considered in settings for which the restriction of a single correlation parameter characterizing the entire scaling matrix $\tilde{\Sigma}_t$ might be too restrictive empirically. For example, we might want to allow for different dependence between financial firms in stressed and non-stressed countries in the context of the euro area sovereign debt crisis. For two blocks containing $N_1$ and $N_2$ firms, respectively, the 2-block equicorrelation specification is given by

$$\tilde{\Sigma}_t = \begin{pmatrix} (1 - \rho^2_{1,t})I_{N_1} & 0 \\ 0 & (1 - \rho^2_{2,t})I_{N_2} \end{pmatrix} + \begin{pmatrix} \rho_{1,t}\ell_{N_1} & \rho_{1,t}\ell_{N_2} \\ \rho_{2,t}\ell_{N_1} & \rho_{2,t}\ell_{N_2} \end{pmatrix} \cdot \begin{pmatrix} \rho_{1,t}\ell'_{N_1} & 0 \\ 0 & \rho_{2,t}\ell'_{N_2} \end{pmatrix}. \quad (13)$$

The 2-block equicorrelation structure (13) differs from the setup in Engle and Kelly (2012) in that there is a direct relation between the equicorrelation in the off-diagonal blocks and the diagonal blocks. The main advantage of this specification is that conditional on $\varsigma_t$ equation (13) preserves the Vasicek (1987) single factor credit risk structure of $y_t$. In particular, specification (12) depends on two common factors only, namely $\kappa_t$ and $\varsigma_t$, and conditionally on $\varsigma_t$ the model is linear. We use this feature extensively to fast and reliably compute joint and conditional risk measures in Section 3 in settings where standard simulation methods quickly become inefficient.\footnote{We demonstrate in Section 4 that the tail risk measurements obtained from applying (10) or (13) are close to those based on a full correlation matrix analysis.}

If $\tilde{\Sigma}_t$ is given by (13) with $\rho_{j,t} = (1 + \exp(-f_{j,t}))^{-1}$ for $j = 1, 2$, then the time varying factor $f_t = (f_{1,t}, f_{2,t})' \in \mathbb{R}^{2 \times 1}$ follows (9), with

$$\Psi_t = \frac{\partial \text{vec}(\tilde{\Sigma}_t)'}{\partial f_t} = \frac{\partial \text{vec}(\tilde{\Sigma}_t)'}{\partial \rho_t} \frac{\partial \rho_t'}{\partial f_t}, \quad (14)$$

$$\frac{\partial \rho_t'}{\partial f_t} = \begin{pmatrix} \exp(-f_{1,t}) \frac{(1+\exp(-f_{1,t}))^2}{(1+\exp(-f_{1,t}))} & 0 \\ 0 & \exp(-f_{2,t}) \frac{(1+\exp(-f_{2,t}))^2}{(1+\exp(-f_{2,t}))} \end{pmatrix},$$

$$\frac{\partial \text{vec}(\tilde{\Sigma}_t)'}{\partial \rho_t} = \begin{pmatrix} \text{vec} \begin{pmatrix} I_{N_1} & 0 \\ 0 & 0 \end{pmatrix}, \text{vec} \begin{pmatrix} 0 & 0 \\ 0 & I_{N_2} \end{pmatrix} \end{pmatrix} \cdot \begin{pmatrix} -2\rho_{1,t} & 0 \\ 0 & -2\rho_{2,t} \end{pmatrix}$$

$$+ \begin{pmatrix} \rho_{1,t}\ell_{N_1} \\ \rho_{2,t}\ell_{N_2} \end{pmatrix} \otimes I_N + I_N \otimes \begin{pmatrix} \rho_{1,t}\ell_{N_1} \\ \rho_{2,t}\ell_{N_2} \end{pmatrix} \cdot \begin{pmatrix} \ell_{N_1} \\ 0 \end{pmatrix} \cdot \begin{pmatrix} \ell_{N_2} \\ 0 \end{pmatrix}.$$
where \( \rho_t = (\rho_{1,t}, \rho_{2,t})' \) and \( N = N_1 + N_2. \)

The \( m \)-block equicorrelation structure is a straightforward generalization of the 2-block case (14). Rather than providing the (lengthy) expressions in the main text, we refer to our Web Appendix for the precise formulations. These formulations are used in our empirical analysis in Section 4, where we also consider a 3-block equicorrelation specification.

### 2.4 Parameter estimation

We can estimate the static parameters \( \theta \) of the dynamic GHST model by standard maximum likelihood procedures. Estimation is straightforward as the likelihood function is known in closed form using a standard prediction error decomposition. Deriving the asymptotic behavior for time varying parameter models with GAS dynamics is non-trivial. We refer to Blasques et al. (2012, 2014) for details.

We split the estimation problem into two parts by adopting a copula perspective. As a result, the number of parameters that need to be estimated in each step is reduced substantially. In addition, the copula perspective has the advantage that we can add more flexibility to modeling the marginal distributions. For example, when working with a multivariate GHST density, all marginal distributions must have the same kurtosis parameter \( \nu \). By adopting a copula perspective, we can relax this restriction considerably.

The two stages of the estimation process can be summarized as follows. In a first step, we estimate univariate dynamic GHST models using the equity returns for each firm \( i \); see the Web Appendix for details. Based on the estimated univariate models with parameters \( \tilde{\mu}_t, \tilde{\sigma}_t, \gamma_t, \) and \( \nu_t \), we transform the observations into their probability integral transforms \( u_{it} \in [0, 1] \). In the second step, we estimate the matrix \( \tilde{\Sigma}_t = \tilde{R}_t \) as parameterized in Section 2.3 using the probability integral transforms \( u_{it} \) constructed in the first step. The GHST copula parameters are \( 0, \tilde{R}_t, \gamma \cdot (1, \ldots, 1)' \) for \( \gamma \in \mathbb{R} \), and \( \nu \), respectively. The static parameter vector \( \theta \) includes \( \gamma, \nu, \tilde{\omega}, A_j, \) and \( B_j \) of the dynamic equation (7).

---

\(^4\) We can introduce further flexibility to the model by extending the support of \( \rho_{1,t} \) and \( \rho_{2,t} \) from \((0, 1)\) to \((-1, 1)\) by defining \( \rho_{j,t} = (\exp(f_{j,t}) - 1)/(\exp(f_{j,t}) + 1) \) for \( j = 1, 2 \). The advantage of this extension is that the correlations between blocks can now become negative, whereas the within block correlation remains positive. This extension is not needed for our empirical application in Section 4, however.
3 Joint and conditional risk measures

This section defines joint and conditional risk measures and demonstrates how to compute these efficiently and reliably based on the application of a conditional law of large numbers (cLLN). The use of a cLLN in the credit risk context has been popularized by Vasicek (1987) and studied further in for example Gordy (2000, 2003) and Lucas et al. (2001, 2003).

Conceptually, the simplest way to compute joint and conditional default probabilities is based on Monte Carlo simulations of firms’ asset values. For example, one can generate many paths for the joint evolution of \((y_{1t}, \ldots, y_{Nt})\) and check how many simulations lie in a joint distress region of the type \(\{y_t \mid y_{jt} < y^*_{jt}, \forall j \in J\}\) for some set of firms \(J \subset \{1, 2, \ldots, N\}\), where \(y^*_{jt}\) denotes the default threshold for firm \(j\) at time \(t\). Such a simulation-based approach quickly becomes inefficient if the cross sectional dimension of the data and the number of firms considered in the set \(J\) grow large: because marginal default probabilities are typically small, we need a large number of simulations to obtain a sufficient number of realizations of joint defaults, particularly if 3, 4, or even more joint defaults are considered. A partial remedy could be to use simulations based on importance sampling methods as in Glassermann and Li (2005), but the computational burden would remain high compared to the simple semi-analytic approach proposed in this section.

The semi-analytic cLLN approximation we propose here is based on the observation that (12) is a non-linear 2-factor model. More specifically, conditional on the common factor \(\varsigma_t\) model (12) simplifies to a (heterogeneous) Gaussian 1-factor model as in Vasicek (1987). This holds even if we replace \(\rho_t\) in (12) by \(\rho_{it}\) using the block-equicorrelation structure from Section 2. We exploit this feature to obtain reliable alternative risk measures that can be evaluated semi-analytically.

We define our joint tail risk measure (JRM) as the time-varying probability that a certain fraction of firms defaults over a pre-specified period. Let \(D_{N,t}\) denote the fraction of firms defaulting over period \(t\), e.g., \(D_{N,t} = 5\%\), with

\[
D_{N,t} = \frac{1}{N} \sum_{i=1}^{N} 1\{y_{it} < y^*_{it}\}.
\]
Since the indicators $1\{y_{it} < y_{it}^*\}$ are conditionally independent given $\kappa_t$ and $\varsigma_t$, we can apply a conditional law of large numbers to obtain

$$D_{N,t} \approx \frac{1}{N} \sum_{i=1}^{N} E[1\{y_{it} < y_{it}^*\} | \kappa_t, \varsigma_t] = \frac{1}{N} \sum_{i=1}^{N} P[y_{it} < y_{it}^* | \kappa_t, \varsigma_t] \equiv C_{N,t},$$

for large $N$, as under standard regularity conditions $|D_{N,t} - C_{N,t}| \overset{a.s.}{\to} 0$ for $N \to \infty$; see for example Vasicek (1987), Gordy (2000, 2003), and Lucas et al. (2001). Note that

$$P[y_{it} < y_{it}^* | \kappa_t, \varsigma_t] = \Phi\left(\frac{y_{it}^* - (\varsigma_t - \mu_\varsigma) \gamma_i - \sqrt{\varsigma_t \rho_t \kappa_t}}{\sqrt{\varsigma_t (1 - \rho_t^2)}}\right),$$

where $\Phi(\cdot)$ denotes the cumulative standard normal distribution. Also note that $C_{N,t}$ is a function of the random variables $\kappa_t$ and $\varsigma_t$ only, and not of $u_{it}$ in (12). We now define the joint tail risk measure as

$$P(D_{N,t} > \bar{c}) \approx P(C_{N,t} > \bar{c}) = P(C_{N,t}(\kappa_t, \varsigma_t) > \bar{c}) \equiv p_t,$$

i.e., we approximate the probability that the fraction of credit portfolio defaults $D_{N,t}$ exceeds the threshold $\bar{c} \in [0, 1]$ by the quantity $p_t$. Following Vasicek (1987) and using the 1-factor structure of (12) for given $\varsigma_t$, we note that $C_{N,t}$ is monotonically decreasing in $\kappa_t$ for $\rho_{i,t} > 0$. This is intuitive: an increase in $\kappa_t$ (for instance due to improved business cycle conditions) implies less defaults in the portfolio and vice versa. We exploit this to efficiently compute unique threshold levels $\kappa_{N,t}^*\left(\bar{c}, \varsigma^{(g)}\right)$ for a number of grid points $\varsigma^{(g)}$, $g = 1, \ldots, G$. This can be done by solving the equations $C_{N,t}(\kappa_{N,t}^*\left(\bar{c}, \varsigma^{(g)}\right), \varsigma^{(g)}) = \bar{c}$ numerically for the threshold values $\kappa_{N,t}^*\left(\bar{c}, \varsigma^{(g)}\right)$ for each grid point $\varsigma^{(g)}$ and time $t$. Given a grid of threshold values, we can then use standard numerical integration techniques to efficiently compute the joint default probability

$$p_t = P(C_{N,t} > \bar{c}) = \int P(\kappa_t < \kappa_{N,t}^*\left(\bar{c}, \varsigma_t\right)) p(\varsigma_t) d\varsigma_t.$$

15
Our second measure is a conditional tail risk measure (CRM). Let
\[ D_{N-1,t}^{(-i)} = \frac{1}{N-1} \sum_{j \neq i} 1[y_{jt} < y_{jt}^* | \kappa_t, \varsigma_t] \approx \frac{1}{N-1} \sum_{j \neq i} P[y_{jt} < y_{jt}^* | \kappa_t, \varsigma_t] = C_{N-1,t}^{(-i)}, \]
where \( D_{N-1,t}^{(-i)} \) is the fraction of defaulted companies excluding firm \( i \), which we approximate using the cLLN by \( C_{N-1,t}^{(-i)} \). We define the CRM as the probability of \( C_{N-1,t}^{(-i)} \) exceeding \( \bar{c}^{(-i)} \) conditional on the default of firm \( i \), i.e.,
\[
P(C_{N-1,t}^{(-i)} > \bar{c}^{(-i)} \bigm| y_{it} < y_{it}^*) = p_{it}^{-1} \cdot P(C_{N-1,t}^{(-i)} > \bar{c}^{(-i)}, y_{it} < y_{it}^*)
= p_{it}^{-1} \cdot \int P(\kappa_t < \kappa_{N-1,t}^{(-i)}(\bar{c}^{(-i)}, \varsigma_t), y_{it} < y_{it}^* | \varsigma_t) p(\varsigma_t) d\varsigma_t
= p_{it}^{-1} \cdot \int \Phi_2(\kappa_{N-1,t}^{(-i)}(\bar{c}^{(-i)}, \varsigma_t), y_{it}^{**}(\varsigma_t) ; \rho_t) p(\varsigma_t) d\varsigma_t, \tag{20}\]
where \( y_{it}^{**}(\varsigma_t) = (y_{it}^* - (\varsigma_t - \mu)\gamma_i)/\sqrt{\varsigma_t} \), and \( \Phi_2(\cdot, \cdot, \rho_t) \) denotes the cumulative distribution function of the bivariate normal with standard normal marginals and correlation parameter \( \rho_t \). To obtain the last equality in (20), note that the GHST distribution becomes Gaussian conditional on \( \varsigma_t \). The conditional probability (20) is a time-varying higher-frequency extension of the multivariate extreme spillovers measure of Hartmann et al. (2004, 2007).

Both the joint probability (19) and the conditional probability (20) can are fast to compute using simple 1-dimensional numerical integration techniques.\(^5\) In addition, the model is easily extended to fit the \( m \)-block equicorrelation structure explained in Section 2.3. The fact that the \( \rho_{i,t} \) parameters are different between blocks does not distort the one-factor structure à la Vasicek (1987) of the current set-up, and all computations remain of similar structure and speed.

\(^5\)It is straightforward to add exposure weights \( e_i \) to the definition of \( D_{N,t} \) in (15). The computations in that case remain equally efficient. If the exposures are very unevenly distributed, however, the approximation error of the cLLN in (16) might increase. To mitigate such an effect, one could try to implement a second-order expansion using a conditional central limit theorem rather than a cLLN only.
4 Euro area financial sector joint tail risk

We apply our model to a high-dimensional dataset of $N = 73$ euro area financial firms. Before presenting the results for the entire dataset, we first present a preliminary analysis for a subsample of $N = 10$ large firms headquartered in a different EU countries. This allows us to benchmark the results for the dynamic GHST block equicorrelation model against a model with a fully specified time varying correlation matrix and to investigate the sensitivity of joint and conditional tail risk measures to the (block) equicorrelation assumption.

4.1 Equity and EDF data

Our equity data come from Bloomberg. We use 73 listed financial firms are located in 11 euro area countries: Austria (AT), Belgium (BE), Germany (DE), Spain (ES), Finland (FI), France (FR), Greece (GR), Ireland (IE), Italy (IT), the Netherlands (NL), and Portugal (PT). Firms were selected if (i) they were financial firms headquartered in the euro area, and (ii) were listed as of 2011Q1 as a component of the STOXX Europe 600 index. For each firm, we construct recursively demeaned weekly equity returns. The sample comprises large commercial banks as well as large financial non-banks such as insurers and investment companies. The total panel covers 762 weeks from January 1999 to August 2013. The panel is unbalanced in that some data are missing in the first part of the sample. We assume that this is not related to the volatility dynamics or the credit risk mechanism in the data. The scores in equation (9) automatically correct for the unbalancedness of the data.

For the marginal default probabilities $p_{it}$, we use one year ahead expected default frequencies (EDF) obtained from Moody’s Analytics. EDFs are widely-used measures of time-varying one-year-ahead marginal default probabilities (see Duffie et al., 2007). We do not use the EDFs to estimate the dependence structure. Rather, we only use the EDFs to calibrate the model each week to current market perceptions of one-year-ahead marginal default risk conditions for each firm in the sample. The copula as estimated from the equity data subsequently takes care of the dependence structure when computing the joint and conditional tail risk measures.
4.2 Small-sample study of 10 banking groups

Before presenting the results for the full sample of $N = 73$ institutions, we first study a geographically diversified sub-sample of ten large financial firms from ten euro area countries. We do so to study how the equicorrelation assumption may affect our joint and conditional risk measures compared to a model with an unrestricted correlation structure. The subsample contains no missing observations and consists of Erste Bank Group (AT), Dexia (BE), Deutsche Bank (DE), Santander (ES), BNP Paribas (FR), National Bank of Greece (GR), Bank of Ireland (IE), UniCredito (IT), ING (NL), and Banco Comercial Portugues (PT). We distinguish firms across different countries given the interdependence of bank and sovereign risk as an important feature of the euro area sovereign debt crisis; see for example ECB (2012, 2014).

4.2.1 Dependence modeling

This section compares correlation estimates across four different dependence models. Since we use a copula approach, the models share the same structure for the univariate volatilities. The descriptive statistics in the Web Appendix reveal that the return data are significantly negatively skewed and fat-tailed. We therefore use the dynamic GHST model for the marginals.

We consider three score-driven equicorrelation specifications with 1, 2, and 3 blocks, respectively. The 2-block model distinguishes firms in countries that experienced pronounced stress during the sovereign debt crisis (Greece, Ireland, Italy, Portugal, Spain) and firms headquartered in non-stressed countries; see Eser and Schwaab (2015) for a similar grouping of countries. The 3-block model further distinguishes firms in smaller stressed countries that entered the sovereign debt crisis earlier (Greece, Portugal, Ireland), and larger stressed countries that entered the sovereign debt crisis at a relatively later time (Spain and Italy); see ECB (2014). As a benchmark, we consider model with a full correlation matrix with DCC dynamics as in Christoffersen et al. (2014), which we estimate by composite likelihood methods. As in the block-equicorrelation specifications, the DCC full correlation matrix
Table 1: Parameter estimates for the copula models

Parameter estimates for multivariate GAS-GHST models for financial firms’ equity returns. The left-hand and right-hand blocks refer to the copula models for \( N = 10 \) and \( N = 73 \) firms, respectively. Univariate GAS-GHST models are used to model the marginal volatilities. The 2-block model distinguishes between firms from stressed countries, i.e., Greece, Portugal, Spain, Italy, Ireland, and firms from the remaining non-stressed countries in the euro area, i.e., Austria, Belgium, France, Germany, and the Netherlands. The 3-block model further distinguishes between financial firms from smaller stressed countries (Greece, Portugal, Ireland), larger stressed countries (Spain, Italy), and non-stressed countries. Standard errors are in parentheses.

<table>
<thead>
<tr>
<th>A</th>
<th>GAS-BEq[1] 0.202 (0.058)</th>
<th>GAS-BEq[2] 0.111 (0.022)</th>
<th>GAS-BEq[3] 0.099 (0.015)</th>
<th>0.021 (0.003)</th>
<th>GAS-BEq[1] 0.544 (0.057)</th>
<th>GAS-BEq[1]-t 0.261 (0.031)</th>
<th>GAS-BEq[3] 0.165 (0.020)</th>
</tr>
</thead>
<tbody>
<tr>
<td>B</td>
<td>0.983 (0.015)</td>
<td>0.991 (0.007)</td>
<td>0.993 (0.005)</td>
<td>0.966 (0.003)</td>
<td>0.984 (0.011)</td>
<td>0.980 (0.011)</td>
<td>0.992 (0.005)</td>
</tr>
<tr>
<td>( \omega_1 )</td>
<td>0.569 (0.210)</td>
<td>0.764 (0.284)</td>
<td>0.713 (0.293)</td>
<td>0.179 (0.236)</td>
<td>0.333 (0.171)</td>
<td>0.167 (0.193)</td>
<td></td>
</tr>
<tr>
<td>( \omega_2 )</td>
<td>0.201 (0.318)</td>
<td>0.870 (0.451)</td>
<td>0.459 (0.191)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \omega_3 )</td>
<td>-0.341 (0.446)</td>
<td>-0.387 (0.335)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \nu )</td>
<td>17.106 (2.012)</td>
<td>17.683 (2.135)</td>
<td>18.571 (2.705)</td>
<td>14.784 (1.791)</td>
<td>32.534 (1.809)</td>
<td>34.802 (2.152)</td>
<td>31.359 (2.673)</td>
</tr>
<tr>
<td>( \gamma_1 )</td>
<td>-0.230 (0.092)</td>
<td>-0.237 (0.106)</td>
<td>-0.250 (0.117)</td>
<td>-0.294 (0.058)</td>
<td>-0.586 (0.054)</td>
<td>-0.590 (0.098)</td>
<td></td>
</tr>
<tr>
<td>( \gamma_2 )</td>
<td>-0.307 (0.101)</td>
<td>-0.272 (0.132)</td>
<td>-0.433 (0.068)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \gamma_3 )</td>
<td>-0.352 (0.126)</td>
<td>-0.574 (0.087)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

The first columns in Table 1 report parameter estimates and model selection criteria for the 1, 2, and 3-block equicorrelation models for \( N = 10 \). The correlation dynamics are highly persistent in all specifications given the high values of \( B \), or of \( A + B \) in case of the DCC specification. The unconditional correlation levels as captured by the parameters \( \omega_i \) are higher for firms in non-stressed countries. For \( N = 10 \), both the degrees of freedom parameter and the common skewness parameter of the GHST copula have similar sign and size for the block equicorrelation models (BEq[m]) and the DCC specification.
Figure 1 plots the estimated dynamic correlations. The top left panel plots the single block equicorrelation along with the average pairwise correlations of each firm with the other (nine) firms. The latter correlation estimates are based on the full-correlation model with 45 time varying parameters. The correlation estimates reveal a pronounced commonality in the correlation dynamics. This is intuitive, as we model firms from the same industry which operate in a single currency area and are subject to a similar regulatory requirements. All correlations tend to increase over the sample period, possibly reflecting gradual financial integration and economic convergence in the euro area after the inception of the euro in 1999. All correlations remain elevated during the global financial crisis from 2008–2010, and peak at a time when Greece, Ireland, and Portugal needed the assistance of third parties such as the EU and the IMF in mid-2010, see ECB (2014). Such shared correlation dynamics can be captured simply and conveniently by block equicorrelation structures.

We can capture a larger share of the cross-sectional dispersion in correlations when we allow for multiple blocks. The top right panel of Figure 1 plots the dynamic correlation estimates for two groups. The first group contains the Bank of Ireland, Banco Comercial Portugues, Santander, UniCredit and the National Bank of Greece. The second group includes BNP Paribas, Deutsche Bank, Dexia, Erste Bank Group, and ING. The overall dependence dynamics are similar. The bottom left panel plots the correlation estimates for the 3-block model, which allows for further cross-sectional dispersion in correlation estimates. Finally, the bottom right panel compares the average correlations across models. In addition, we also provide average correlations estimated from a 52-week rolling window. The average correlations are similar across all models. Only relatively minor deviations are observed. We conclude that despite its restrictive nature the equicorrelation model reliably estimates the salient trends in average correlation dynamics.

Although the block equicorrelation models work well in capturing the average correla-

\footnote{A principal components analysis of the 45 correlation pairs from the full-correlation model suggest that the first three components explain 55.6%, 16%, and 8.5% of the total variation, respectively. The first two factors therefore explain approximately 72% of the total variation in correlations.}

\footnote{Perhaps surprisingly, the correlation between firms from non-stressed countries lies above that of firms in stressed countries, also before the financial and sovereign debt crisis starting 2008. It is therefore probably related to different degrees of financial integration, rather than to shared exposure to heightened market turmoil in a crisis.}
Figure 1: Filtered correlation estimates for $N = 10$ firms

The top left panel plots the correlation estimates based on the BEq[1] model, along with the average pairwise correlation of each firm with the other nine firms. The latter are based on the specification with full correlation matrix with 45 time varying parameters and DCC dynamics. The top right and bottom left panels plot the block equicorrelation estimates based on the 2-block (BEq[2]) and 3-block (BEq[3]) model specification, respectively. The bottom right panel compares the average correlation estimates from the 1-block, 2-block, and 3-block models to the average (over 45 pairs) correlation estimates from a GHST full-correlation DCC model, as well as the average correlation based on a 52-week rolling window.

tions, a substantial share of the cross-sectional dispersion in correlations may be lost when using block equicorrelation structures; see the top left panel of Figure 1. This may or may not matter when evaluating joint and conditional credit risk measures. To investigate how much cross-sectional dispersion in correlations is lost, the left panel of Figure 2 plots $R^2$ statistics that correspond to repeated cross-sectional regressions of 45 correlation pairs from the GHST full-correlation-matrix model on a constant and the corresponding correlation estimates from a 2-block and 3-block equicorrelation model, respectively. By construction, the 1-block equicorrelation model is unable to account for any cross-sectional variation in correlations, as it collapses the latter to a single number (i.e., a regression on a constant). For the current sample, the 2-block and 3-block equicorrelation models are able to account
Figure 2: R-squared statistics

The left panel plots $R^2$ statistics corresponding to repeated (over time) cross-sectional regressions of 45 correlation pairs from the $N = 10$ GHST full-correlation-matrix model on a constant and the corresponding correlation estimates from a 2-block, and 3-block equicorrelation model. The right panel plots Kernel density estimates and a histogram of 45 $R^2$ statistics corresponding to time series regressions with $T = 762$ observations each of correlations from a GHST full-correlation-matrix model on the corresponding correlation estimates from a 1-block, 2-block, and 3-block equicorrelation model, respectively. The histogram refers to $R^2$s from the 3-block equicorrelation model.

for approximately 20% and 50% of the cross-sectional dispersion in correlations.

The right panel of Figure 2 plots kernel density estimates and a histogram of time-series $R^2$ statistics. These correspond to a histogram of 45 time series regressions, with $T = 762$ weekly observations each, of GHST full-correlation matrix model correlations on the corresponding estimates from a 1-block, 2-block, and 3-block equicorrelation specification, respectively. The 1-block, 2-block, and 3-block equicorrelation models are able to account for approximately 30-50%, 40-60%, and 50-70% of the time series variation in the full correlation estimates, respectively. Both findings confirm that a large part of the variation in correlations can be captured by the simplified equicorrelation assumption. In the next section we highlight that the percentage of variation captured by the BEq[$m$] models appears adequate for reliable estimates of our systemic risk measures.

4.2.2 Joint and conditional probabilities

This section compares joint and conditional default probabilities as defined in Section 3 and as implied by different copula model specifications. We argue that the cLLN works well even when the data dimension is very small. This is because it eliminates a source of risk
(idiosyncratic risk) that does not matter in the tail of the credit loss distribution. Only \( \kappa_t \) and \( \varsigma_t \) are common to all firms and drive extreme default clustering. In addition, block equicorrelation and full correlation models lead to approximately similar patterns for joint and (to a lesser degree) also for conditional tail probability estimates.

We compare the full-correlation specification with a 1-block and 3-block copula model. In addition, we assess the adequacy of the cLLN approximation by comparing it to a simulation-based approach to compute joint and conditional default probabilities. The default thresholds are obtained by inverting the GHST distribution function at the observed EDF levels. For the simulation-based computations, we use 500,000 simulations, at each time \( t \), where we save the number and the identities of the defaulted firms in each simulation.

The top left and right panels in Figure 3 refer to the probability of observing three or more defaults (out of ten firms) over a one-year-ahead horizon. The left panel compares the 1-block specification with the full-correlation matrix outcome; the right panel considers the 3-block specification. Risk measures are either simulated or computed semi-analytically based on equation (19). As an important takeaway, the joint default probabilities are similar in each of the six cases. The losses from moving from a full correlation matrix to a much simpler equicorrelation structure are generally small. Also the losses from applying the cLLN approximation vis-à-vis the simulation approach are small in these cases, even though \( N = 10 \) is far from infinity.

The bottom panels in Figure 3 plot the conditional probability of 2 or more (out of 9 possible) credit events given a credit event for a given financial firm, averaged over all ten firms. Again, the left and right panels refer to the 1-block and 3-block model, respectively. In the 1-block case, the loss in fit from moving from a full-correlation-matrix model to the equicorrelation model is more pronounced. This is in line with the \( R^2 \) results shown in Figure 2. The cLLN approximation, however, works relatively well. In the 3-block case, the respective loss in fit is less pronounced. Here, however, the cLLN approximation appears to work slightly less well. This reveals an interesting trade-off when choosing the number of blocks. On the one hand, increasing the number of blocks can help in capturing more of the cross-sectional dispersion in correlations and increases the fit in the time dimension,
Figure 3: Joint and conditional probabilities

The two top panels plot the joint default probability (19) of three or more financial firm defaults out of ten. The two bottom panels plot the (average) conditional probability (20), which is the probability of 2 or more (out of 9 possible) defaults given a default of a given financial firm, averaged over all ten firms. The probabilities are either computed using 500,000 simulations at each point in time $t$, or alternatively using the cLLN approximation as discussed in Section 3 for block equicorrelation models. The left (top and bottom) panels are based on the estimated 1-block equicorrelation model, while the right (top and bottom) panels are based on the estimated 3-block equicorrelation model.

see Figure 2. On the other hand, increasing the number of blocks also means less firms in each block, which in turn implies that the cLLN approximation works less well within each block.\footnote{Recall that the 3-block model contains 5, 3, and 2 firms in each block. The cLLN hardly applies in the latter two cases, particularly given the low loading to the common $\kappa_t$ for the stressed small countries shown in Figure 1.} Finally, the simulation based approach can suffer from sizable simulation noise in non-crisis periods (when marginal default probabilities are low, see bottom panels before 2007). No such problems are encountered for the cLLN based approximation.
4.3 All 73 euro area financial firms

Given the encouraging preliminary results for the subsample of $N = 10$ institutions, we now turn to the joint and conditional tail risk measures based on the full panel of 73 large financial sector firms. The sample contains commercial banks as well as financial non-banks such as insurers and investment companies. All are listed at a stock exchange.\(^9\) Based on descriptive statistics as presented in the Web Appendix, univariate GHST models seem appropriate given the skewness and kurtosis features of the equity return data.

4.3.1 Dependence modeling

Figure 4 plots the correlation estimates. For parameter estimates, we refer to Table 1. We use the same model specifications as described in Section 4.2, but now estimated on the full sample of $N = 73$ firms. The equicorrelation estimates (top left) range from low values of approximately 0.1 in 2000 to values as high as 0.6 towards the start of the sovereign debt crisis in 2010. Using the two block structure (top right), there appear to be little differences between financial firms from stressed versus non-stressed countries, except during the peak of the sovereign debt crisis in 2011–12. The bottom left panel of Figure 4 introduces further heterogeneity by modeling the dynamic correlation for firms from Greece, Ireland, and Portugal on the one hand, and Italy and Spain on the other hand. Again, the rationale for this grouping is that Ireland, Greece, and Portugal entered the euro area sovereign debt crisis earlier, and were relatively more stressed, compared with Spain and Italy. This distinction may matter for our inference on financial sector tail risks around 2012. We obtain a lower correlation level among financial institutions from the first group of smaller stressed countries. The correlation rises after the financial crisis up to the start of 2010, and then decreases to pre-crisis levels around 2012. By contrast, the correlation for financial firms from Italy and Spain starts to rise earlier, peaks higher, and remains high until the end of

\(^9\)Freezing the set of firms as the constituents of a broad based equity index in 2011Q1 means that we may underestimate total euro area financial sector risk prior to this date (due to sample selection; weaker firms may have dropped out of the index by then). While this concern should be kept in mind, it is unlikely to be large, as most financial institutions under stress during the financial crisis continued to operate, also due to substantial government aid and the extension of public sector guarantees. Sample selection is no issue after 2011Q1.
the sample. If we consider the average correlations across all \(73(73 - 1)/2 = 2628\) pairs in the bottom right panel, the picture emerging from all three model specifications is similar.

Our equity panel dataset is characterized by a substantial amount of common variation across correlations. A principle components analysis (PCA) of 52-week rolling window correlations suggest that the first three components explain approximately 43%, 22% and 12% of the total variation in pair-wise correlations, respectively. This suggests that a 3-block model specification — which allows for 6 different correlations at each point in time — should capture a substantial share (approximately 60-75%) of the common time series variation of all 2628 full correlations. This is the case. We refer to the Web Appendix for goodness-of-fit statistics from regressing full correlation estimates on BEq\([m]\)-estimates. Full correlation estimates are obtained from either rolling window correlations or based on the maximization of a composite likelihood as in Christoffersen et al. (2012).

### 4.3.2 Joint and conditional tail risk

This section presents our joint and conditional tail risk estimates for financial sector firms in the euro area. For the joint default probability, we consider the probability that \(\bar{c} = 7/73 \approx 10\%\) or more of currently active financial institutions experience a credit event over the next 12 months. The probability of such widespread and massive failure should typically be very small during non-crisis times. We plot the result in the upper panel in Figure 5.

Different copula specifications give strikingly similar results. As the joint probability moves relatively little before 2008, we only plot it over the period 2006–2013. For a comparison of joint and conditional risk outcomes for a variety of GHST and symmetric-t copula specifications we refer to the Web Appendix.\(^{10}\) The joint probability moves sharply upwards after the bankruptcy of Lehman Brothers in September 2008, and reaches a first peak during the Irish sovereign debt crisis in the Spring of 2009. It remains approximately constant thereafter until the onset of the sovereign debt crisis in early 2010. It reaches a first peak in late 2011, followed by a second and final peak mid-2012.

\(^{10}\)The likelihood of the single block equicorrelation model drops by approximately \(10873.8 - 10922.2 = 48.4\) points if the restriction of symmetry is imposed; see Table 1. This is statistically significant.
The vertical lines in the top panel of Figure 5 indicate a number of relevant policy announcements. In late 2011, the announcement and implementation of two 3-year Long Term Refinancing Operations (3y-LTROs, allotted in December 2011 and February 2012) had a visible but temporary impact on financial sector tail risk. In particular, the announcement and allotment of two 3y-LTRO in December 2011 and February 2012 appear to have lowered financial sector joint tail risk, temporarily, for a few months. In the first half of 2012, financial sector joint tail risk picked up again and rose to unprecedented levels until July 2012. The three vertical lines from July to September 2012, together with the time variation in the joint default probability, strongly suggest that three events — a speech by the ECB president in London to do ‘whatever it takes’ to save the euro on 26 July 2012, the announcement of

See ECB (2011) for the official press release and monetary policy objectives.
Figure 5: Joint and average conditional tail probabilities

The top panel plots the joint tail probability that 7 or more financial firms (\( \bar{c} = 7/73 \approx 10\% \)) experience a credit event over a 12 months ahead horizon, at any time \( t \). The bottom panel plots the average conditional risk measure, the average (over all 73 financial firms) conditional probability that \( \bar{c} = 7/73 \approx 10\% \) or more financial firms default given a default for a given firm \( i \). In each case, computations are based on a GAS-BEq[1], GAS-BEq[3], and DCC-CL full correlation model, respectively. Computations rely on the conditional law of large numbers approximation for the GHST BEq\( m \) models, and on simulation for the DCC-composite likelihood model.

The vertical lines in the top panel indicate the following events: (a) the allotments of two three-year long term refinancing operations by the ECB on 21 December 2011 and 29 February 2012, (b) a speech by the ECB president (“whatever it takes”) on 26 July 2012, (c) the OMT announcement on 2 August 2012, and (d) the announcement of the OMT technical details on 6 September 2012. The vertical line in the lower panel indicates the Lehman Brothers bankruptcy on 15 September 2008.
the ECB’s Outright Monetary Transactions program on 2 August 2012, and especially the disclosure of the OMT details on 6 September 2012 — collectively ended the most acute phase of extreme financial sector joint tail risks. The joint default probability decreases sharply, all the way until the end of the sample. The OMT is a program within which the ECB can make purchases (“outright transactions”) in secondary sovereign bond markets, under certain conditions, of bonds issued by euro area member states. We refer to ECB (2012) and Coeuré (2013) for details. No purchases have yet been undertaken by the ECB within the OMT. Instead, the mere announcement of the program was sufficient to trigger the substantial decline in financial sector joint tail risk.

The bottom panel of Figure 5 presents the average (across institutions) conditional tail probability for the 73 firms in our sample. Again, different copula model specifications give approximately similar results, although there are some discrepancies in the first third of the sample leading up to the financial crisis. Conditional probability estimates are also more sensitive to the choice of skewed or non-skewed copula; see the Web Appendix for details. Also, the DCC based conditional measures appear to be much less responsive to the different historical events. This is in line with the well-known phenomenon that the dynamics of DCC models become relatively flat in high dimensions. The equicorrelation model is much less susceptible to this bias. Though the composite likelihood approach for estimation partly corrects the potential bias in the DCC estimates, Figure 5 illustrates that the equicorrelation model pick up more of the dynamics in the series.

At the start of the sample, there is little evidence of systemic clustering on average with low levels of the CRM between 10%–40%. The conditional probability rises after the Lehman Brothers bankruptcy to levels of approximately 60%, and then to approximately 80% around the peak of the sovereign debt crisis. Such high levels of conditional tail probabilities signal strong interconnectedness among euro area financial institutions. Interestingly, the conditional tail probability is still quite high towards the end of the sample, despite the collapse in joint risk as shown in the top panel. The OMT announcements apparently did not bring down markets’ perceptions of conditional (or contagion) risks in the euro area financial system as a whole to a comparable extent.
As a final result, unconventional monetary policy measures, such as the 3y-LTROs and the OMT announcements, and financial stability tail risk outcomes appear strongly related. This suggests substantial scope for the coordination of monetary, macro-prudential, and bank supervision policies. This is relevant as both monetary policy as well as prudential policies and banking supervision will be carried out jointly within the ECB as of November 2014.

5 Conclusion

We developed a novel modeling framework for estimating joint and conditional tail risk probabilities over time in a financial system that consists of many financial sector firms. For this purpose, we used a copula approach based on the generalized hyperbolic skewed Student’s $t$ (GHST) distribution, endowed with score-driven dynamics. Parsimony and flexibility were traded off by using a dynamic block equicorrelation structure. Using this structure, we were able to implement efficient approximations based on a conditional law of large numbers to compute joint and conditional tail probabilities of multiple defaults for a large set of firms. An application to euro area financial firms from 1999 to 2013 revealed unprecedented joint default risks between 2011–12. We also document the collapse of these joint default risks (but not conditional risks) after a sequence of announcements pertaining to the ECB’s Outright Monetary Transactions program in August and September 2012.

References


34
<table>
<thead>
<tr>
<th>Title</th>
<th>Authors</th>
<th>Date</th>
</tr>
</thead>
<tbody>
<tr>
<td>Estimation of an Adaptive Stock Market Model with Heterogeneous Agents</td>
<td>Henrik Amilon</td>
<td>2005:177</td>
</tr>
<tr>
<td>Some Further Evidence on Interest-Rate Smoothing: The Role of Measurement Errors in the Output Gap</td>
<td>Mikael Apel and Per Jansson</td>
<td>2005:178</td>
</tr>
<tr>
<td>Bayesian Estimation of an Open Economy DSGE Model with Incomplete Pass-Through</td>
<td>Malin Adolfson, Stefan Laséen, Jesper Lindé and Mattias Villani</td>
<td>2005:179</td>
</tr>
<tr>
<td>Are Constant Interest Rate Forecasts Modest Interventions? Evidence from an Estimated Open Economy DSGE Model of the Euro Area</td>
<td>Malin Adolfson, Stefan Laséen, Jesper Lindé and Mattias Villani</td>
<td>2005:180</td>
</tr>
<tr>
<td>Inference in Vector Autoregressive Models with an Informative Prior on the Steady State</td>
<td>Mattias Villani</td>
<td>2005:181</td>
</tr>
<tr>
<td>Bank Mergers, Competition and Liquidity</td>
<td>Elena Carletti, Philipp Hartmann and Giancarlo Spagnolo</td>
<td>2005:182</td>
</tr>
<tr>
<td>Exploring Interactions between Real Activity and the Financial Stance</td>
<td>Tor Jacobson, Jesper Lindé and Kasper Roszbach</td>
<td>2005:184</td>
</tr>
<tr>
<td>Two-Sided Network Effects, Bank Interchange Fees, and the Allocation of Fixed Costs</td>
<td>Mats A. Bergman</td>
<td>2005:185</td>
</tr>
<tr>
<td>Real Exchange Rate and Consumption Fluctuations following Trade Liberalization</td>
<td>Kristian Jönsson</td>
<td>2005:187</td>
</tr>
<tr>
<td>Bayesian Inference of General Linear Restrictions on the Cointegration Space</td>
<td>Mattias Villani</td>
<td>2005:189</td>
</tr>
<tr>
<td>Forecasting Performance of an Open Economy Dynamic Stochastic General Equilibrium Model</td>
<td>Malin Adolfson, Stefan Laséen, Jesper Lindé and Mattias Villani</td>
<td>2005:190</td>
</tr>
<tr>
<td>Forecast Combination and Model Averaging using Predictive Measures</td>
<td>Jana Eklund and Sune Karlsson</td>
<td>2005:191</td>
</tr>
<tr>
<td>A Simultaneous Model of the Swedish Krona, the US Dollar and the Euro</td>
<td>Hans Lindblad and Peter Sellin</td>
<td>2006:193</td>
</tr>
<tr>
<td>Efficient Bayesian Inference for Multiple Change-Point and Mixture Innovation Models</td>
<td>Paolo Giordani and Robert Kohn</td>
<td>2006:196</td>
</tr>
<tr>
<td>Derivation and Estimation of a New Keynesian Phillips Curve in a Small Open Economy</td>
<td>Karolina Holmberg</td>
<td>2006:197</td>
</tr>
<tr>
<td>Monetary Policy and Staggered Wage Bargaining when Prices are Sticky</td>
<td>Mikael Carlsson and Andreas Westermark</td>
<td>2006:199</td>
</tr>
<tr>
<td>The Swedish External Position and the Krona</td>
<td>Philip R. Lane</td>
<td>2006:200</td>
</tr>
<tr>
<td>Title</td>
<td>Year:Page</td>
<td></td>
</tr>
<tr>
<td>----------------------------------------------------------------------</td>
<td>------------</td>
<td></td>
</tr>
<tr>
<td>Re-Evaluating Swedish Membership in EMU: Evidence from an Estimated Model</td>
<td>2008:227</td>
<td></td>
</tr>
<tr>
<td>Expectation Driven Business Cycles with Limited Enforcement</td>
<td>2009:229</td>
<td></td>
</tr>
<tr>
<td>Effects of Organizational Change on Firm Productivity</td>
<td>2009:230</td>
<td></td>
</tr>
<tr>
<td>Monetary Policy Trade-Offs in an Estimated Open-Economy DSGE Model</td>
<td>2009:232</td>
<td></td>
</tr>
<tr>
<td>Flexible Modeling of Conditional Distributions Using Smooth Mixtures of Asymmetric Student T Densities</td>
<td>2009:233</td>
<td></td>
</tr>
<tr>
<td>Forecasting Macroeconomic Time Series with Locally Adaptive Signal Extraction</td>
<td>2009:234</td>
<td></td>
</tr>
<tr>
<td>Evaluating Monetary Policy</td>
<td>2009:235</td>
<td></td>
</tr>
<tr>
<td>Risk Premiums and Macroeconomic Dynamics in a Heterogeneous Agent Model</td>
<td>2010:236</td>
<td></td>
</tr>
<tr>
<td>Picking the Brains of MPC Members</td>
<td>2010:237</td>
<td></td>
</tr>
<tr>
<td>Involuntary Unemployment and the Business Cycle</td>
<td>2010:238</td>
<td></td>
</tr>
<tr>
<td>Housing collateral and the monetary transmission mechanism</td>
<td>2010:239</td>
<td></td>
</tr>
<tr>
<td>The Discursive Dilemma in Monetary Policy</td>
<td>2010:240</td>
<td></td>
</tr>
<tr>
<td>Monetary Regime Change and Business Cycles</td>
<td>2010:241</td>
<td></td>
</tr>
<tr>
<td>Bayesian Inference in Structural Second-Price common Value Auctions</td>
<td>2010:242</td>
<td></td>
</tr>
<tr>
<td>Equilibrium asset prices and the wealth distribution with inattentive consumers</td>
<td>2010:243</td>
<td></td>
</tr>
<tr>
<td>Identifying VARS through Heterogeneity: An Application to Bank Runs</td>
<td>2010:244</td>
<td></td>
</tr>
<tr>
<td>Modeling Conditional Densities Using Finite Smooth Mixtures</td>
<td>2010:245</td>
<td></td>
</tr>
<tr>
<td>The Output Gap, the Labor Wedge, and the Dynamic Behavior of Hours</td>
<td>2010:246</td>
<td></td>
</tr>
<tr>
<td>Density-Conditional Forecasts in Dynamic Multivariate Models</td>
<td>2010:247</td>
<td></td>
</tr>
<tr>
<td>Anticipated Alternative Policy-Rate Paths in Policy Simulations</td>
<td>2010:248</td>
<td></td>
</tr>
<tr>
<td>MOSES: Model of Swedish Economic Studies</td>
<td>2011:249</td>
<td></td>
</tr>
<tr>
<td>The Effects of Endogenous Firm Exit on Business Cycle Dynamics and Optimal Fiscal Policy</td>
<td>2011:250</td>
<td></td>
</tr>
<tr>
<td>Parameter Identification in a Estimated New Keynesian Open Economy Model</td>
<td>2011:251</td>
<td></td>
</tr>
<tr>
<td>Up for count? Central bank words and financial stress</td>
<td>2011:252</td>
<td></td>
</tr>
<tr>
<td>Title</td>
<td>Authors</td>
<td>Year/Volume</td>
</tr>
<tr>
<td>----------------------------------------------------------------------</td>
<td>---------------------------------------------------------------------------------------------</td>
<td>-------------</td>
</tr>
<tr>
<td>Wage Adjustment and Productivity Shocks</td>
<td>Mikael Carlsson, Julián Messina and Oskar Nordström Skans</td>
<td>2011:253</td>
</tr>
<tr>
<td>Taking the Twists into Account: Predicting Firm Bankruptcy Risk with Splines of Financial Ratios</td>
<td>Paolo Giordani, Tor Jacobson, Erik von Schedvin and Mattias Villani</td>
<td>2011:256</td>
</tr>
<tr>
<td>Collateralization, Bank Loan Rates and Monitoring: Evidence from a Natural Experiment</td>
<td>Geraldo Cerqueiro, Steven Ongena and Kasper Roszbach</td>
<td>2012:257</td>
</tr>
<tr>
<td>Labor-Market Frictions and Optimal Inflation</td>
<td>Mikael Carlsson and Andreas Westermark</td>
<td>2012:259</td>
</tr>
<tr>
<td>Output Gaps and Robust Monetary Policy Rules</td>
<td>Roberto M. Billi</td>
<td>2012:260</td>
</tr>
<tr>
<td>The Information Content of Central Bank Minutes</td>
<td>Mikael Apel and Marianna Blix Grimaldi</td>
<td>2012:261</td>
</tr>
<tr>
<td>The Cost of Consumer Payments in Sweden</td>
<td>Björn Segendorf and Thomas Jansson</td>
<td>2012:262</td>
</tr>
<tr>
<td>Trade Credit and the Propagation of Corporate Failure: An Empirical Analysis</td>
<td>Tor Jacobson and Erik von Schedvin</td>
<td>2012:263</td>
</tr>
<tr>
<td>Structural and Cyclical Forces in the Labor Market During the Great Recession: Cross-Country Evidence</td>
<td>Luca Sala, Ulf Söderström and Antonella Trigari</td>
<td>2012:264</td>
</tr>
<tr>
<td>Pension Wealth and Household Savings in Europe: Evidence from SHARELIFE</td>
<td>Rob Alessie, Viola Angelini and Peter van Santen</td>
<td>2013:265</td>
</tr>
<tr>
<td>Long-Term Relationship Bargaining</td>
<td>Andreas Westermark</td>
<td>2013:266</td>
</tr>
<tr>
<td>Using Financial Markets To Estimate the Macro Effects of Monetary Policy: An Impact-Identified FAVAR*</td>
<td>Stefan Pitschner</td>
<td>2013:267</td>
</tr>
<tr>
<td>DYNAMIC MIXTURE-OF-EXPERTS MODELS FOR LONGITUDINAL AND DISCRETE-TIME SURVIVAL DATA</td>
<td>Matias Quiroz and Mattias Villani</td>
<td>2013:268</td>
</tr>
<tr>
<td>Conditional euro area sovereign default risk</td>
<td>André Lucas, Bernd Schwaab and Xin Zhang</td>
<td>2013:269</td>
</tr>
<tr>
<td>Nominal GDP Targeting and the Zero Lower Bound: Should We Abandon Inflation Targeting?*</td>
<td>Roberto M. Billi</td>
<td>2013:270</td>
</tr>
<tr>
<td>Un-truncating VARs*</td>
<td>Ferre De Graeve and Andreas Westermark</td>
<td>2013:271</td>
</tr>
<tr>
<td>Housing Choices and Labor Income Risk</td>
<td>Thomas Jansson</td>
<td>2013:272</td>
</tr>
<tr>
<td>Identifying Fiscal Inflation*</td>
<td>Ferre De Graeve and Virginia Queijo von Heideken</td>
<td>2013:273</td>
</tr>
<tr>
<td>On the Redistributive Effects of Inflation: an International Perspective*</td>
<td>Paola Boel</td>
<td>2013:274</td>
</tr>
<tr>
<td>Business Cycle Implications of Mortgage Spreads*</td>
<td>Karl Walentin</td>
<td>2013:275</td>
</tr>
<tr>
<td>Approximate dynamic programming with post-decision states as a solution method for dynamic economic models</td>
<td>Isaiah Hull</td>
<td>2013:276</td>
</tr>
<tr>
<td>A detrimental feedback loop: deleveraging and adverse selection</td>
<td>Christoph Bertsch</td>
<td>2013:277</td>
</tr>
<tr>
<td>Distortionary Fiscal Policy and Monetary Policy Goals</td>
<td>Klaus Adam and Roberto M. Billi</td>
<td>2013:278</td>
</tr>
<tr>
<td>Predicting the Spread of Financial Innovations: An Epidemiological Approach</td>
<td>Isaiah Hull</td>
<td>2013:279</td>
</tr>
</tbody>
</table>
Firm-Level Evidence of Shifts in the Supply of Credit  
by Karolina Holmberg  
2013:280

Lines of Credit and Investment: Firm-Level Evidence of Real Effects of the Financial Crisis  
by Karolina Holmberg  
2013:281

A wake-up call: information contagion and strategic uncertainty  
by Toni Ahnert and Christoph Bertsch  
2013:282

Debt Dynamics and Monetary Policy: A Note  
by Stefan Laséen and Ingvar Strid  
2013:283

Optimal taxation with home production  
by Conny Olovsson  
2014:284

Incompatible European Partners? Cultural Predispositions and Household Financial Behavior  
by Michael Halilasos, Thomas Jansson and Yigitcan Karabulut  
2014:285

How Subprime Borrowers and Mortgage Brokers Shared the Piecial Behavior  
by Antje Berndt, Burton Hollifield and Patrik Sandás  
2014:286

The Macro-Financial Implications of House Price-Indexed Mortgage Contracts  
by Isaiah Hull  
2014:287

Does Trading Anonymously Enhance Liquidity?  
by Patrick J. Dennis and Patrik Sandás  
2014:288

Systematic bailout guarantees and tacit coordination  
by Christoph Bertsch, Claudio Calcagno and Mark Le Quement  
2014:289

Selection Effects in Producer-Price Setting  
by Mikael Carlsson  
2014:290

Dynamic Demand Adjustment and Exchange Rate Volatility  
by Vesna Corbo  
2014:291

Forward Guidance and Long Term Interest Rates: Inspecting the Mechanism  
by Ferre De Graeve, Pelin Ilbas & Raf Wouters  
2014:292

Firm-Level Shocks and Labor Adjustments  
by Mikael Carlsson, Julián Messina and Oskar Nordström Skans  
2014:293

A wake-up call theory of contagion  
by Toni Ahnert and Christoph Bertsch  
2015:294

Risks in macroeconomic fundamentals and excess bond returns predictability  
by Rafael B. De Rezende  
2015:295

The Importance of Reallocation for Productivity Growth: Evidence from European and US Banking  
by Jaap W.B. Bos and Peter C. van Santen  
2015:296

Amortization Requirements and Household Indebtedness: An Application to Swedish-Style Mortgages  
by Isaiah Hull  
2015:297

Fuel for Economic Growth?  
by Johan Gars and Conny Olovsson  
2015:298

Searching for Information  
by Jungsuk Han and Francesco Sangiorgi  
2015:299

What Broke First? Characterizing Sources of Structural Change Prior to the Great Recession  
by Isaiah Hull  
2015:300

Price Level Targeting and Risk Management  
by Roberto Billi  
2015:301

Central bank policy paths and market forward rates: A simple model  
by Ferre De Graeve and Jens Iversen  
2015:302

Jump-Starting the Euro Area Recovery: Would a Rise in Core Fiscal Spending Help the Periphery?  
by Olivier Blanchard, Christopher J. Erceg and Jesper Lindé  
2015:303

Bringing Financial Stability into Monetary Policy*  
by Eric M. Leeper and James M. Nason  
2015:304