



BANCA D'ITALIA
EUROSISTEMA

Temi di Discussione

(Working Papers)

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by Piergiorgio Alessandri, Antonio Maria Conti and Fabrizio Venditti

June 2017

Number

1121



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ISSN 1594-7939 (print)

ISSN 2281-3950 (online)

Printed by the Printing and Publishing Division of the Bank of Italy

THE FINANCIAL STABILITY DARK SIDE OF MONETARY POLICY

by Piergiorgio Alessandri*, Antonio Maria Conti* and Fabrizio Venditti*

Abstract

Since monetary policy affects risk premiums, and these appear to have a stronger influence on economic activity when they rise than when they fall, temporary monetary expansions may both stimulate the economy and sow the seeds of damaging financial market corrections in the future. We investigate this possibility by using local projection methods to examine the propagation of monetary shocks through US corporate bond markets. We find that, while the transmission of monetary shocks is symmetric, the impact of macroeconomic data releases is asymmetric: spreads are more responsive to bad news. Crucially, these responses precede economic slowdowns rather than directly cause them.

JEL Classification: C32, E32, F34.

Keywords: monetary policy, financial stability, risk premia, macro news, local projections.

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1 Introduction¹

The Great Recession has reinforced the view that financial markets and the real economy are linked in a complex, non-linear way. In particular, although the linkage between credit spreads and economic activity appears to be relatively tenuous in ‘good times’, large spikes in spreads seem capable of causing significant economic contractions in ‘bad times’. At the same time, credit markets have been found to be one of the key linkages in the transmission of monetary policy (Borio and Zhu, 2012; Adrian and Liang, 2014; Buch et al., 2014; Gertler and Karadi, 2015). Piecing the evidence together, Stein (2014) notes that central banks may face a new policy conundrum: if credit spreads have a stronger impact when they tighten, then the benefits of a monetary stimulus that lowers them temporarily might be exceeded by the costs associated to their subsequent reversal, so that all in all monetary policy ends up increasing the volatility rather than the level of economic activity. The existence of such a *financial stability ‘dark side’* to monetary interventions could, once recognized, radically alter their strategy: perhaps *"monetary policy should be less accommodative – [and] tolerate a larger forecast shortfall of the path of the unemployment rate from its full-employment level – when estimates of risk premiums in the bond market are abnormally low"*.² The issue is central to the debate on both the role of monetary policy in the run up to the 2008 financial crisis and to the risks associated to an exit from Quantitative Easing (Bernanke, 2015; Krugman, 2015).

This paper provides an empirical investigation of the dark side argument, i.e. the existence of a sign asymmetry in the transmission of monetary policy shocks to the real economy *via* credit spreads.³ Using the corporate bond spreads constructed by Gilchrist and Zakrajsek (2012) to capture credit conditions, we study the linkage between monetary policy, credit markets and economic activity in the US economy between 1973 and 2012. Our starting point is that, although the reduced-form asymmetry presented by Stein (2014) is both intuitive and empirically plausible, its existence can reflect two structural stories that differ in critical ways. On the one hand, a rise in spreads could indeed have a much stronger impact on economic activity because it forces financially-constrained agents to quickly reduce their leverage (Brunnermeier

¹We wish to thank Tobias Adrian, Giovanni Caggiano, Efram Castelnuovo, Jean Flemming, Marc Giannoni, Giovanni Veronese and participants to the Fourth Workshop on Empirical Macroeconomics in Ghent, the joint BOE, ECB, CEPR and CFM conference on "Credit Dynamics and the Macroeconomy", the 9th CFE Conference, the ESRB 2015 workshop in Vilnius, the 2016 annual conference of the European Economic Association, the 2016 annual conference of the International Association of Applied Econometrics, the Fifth International Conference in memory of Carlo Giannini on "Recent Developments in Econometric Methodologies" and seminars held at Banca d'Italia, Bank of England, Bundesbank, ECB and IMF. All remaining mistakes are the authors' own responsibility. *The views expressed herein are those of the authors and do not necessarily reflect those of the Banca d'Italia or the Eurosystem.*

²Stein (2014), p.1. For further discussion, see also Kocherlakota (2014).

³As explained above, this is a different meaning with respect to the traditional Keynesian asymmetry stating that expansionary monetary policy shocks have smaller effects than contractionary ones, evaluated for example by Ravn and Sola (2004) and, more recently, Tenreiro and Thwaites (2016).

et al., 2009; Guerrieri and Iacoviello, 2013; He and Krishnamurthy, 2013; Brunnermeier and Sannikov, 2014). On the other hand, spreads might move more ahead of a downturn simply because investors are more interested in recessions than expansions and respond more actively to bad news on the economic outlook (Veronesi, 1999; Epstein and Schneider, 2008; Beber and Brandt, 2010). Discriminating among these explanations is obviously important: monetary policy is likely to work asymmetrically in the first case but not in the second. Furthermore, an asymmetry caused by credit constraints may be important from a welfare perspective, whereas a purely predictive one certainly is not. It is also difficult, as the exercise involves identifying monetary policy and macroeconomic “news” shocks and mapping their (potentially non-linear) propagation through credit markets. Our strategy to tackle these challenges relies on two ideas: we identify the shocks exploiting high-frequency data on the reaction of bond markets to monetary surprises and new data releases (Gertler and Karadi, 2015; Faust et al., 2007; Caldara and Herbst, 2016), and we then use local projections to characterize the implications of the shocks in a flexible way (Jordà, 2005).

We find that, in a reduced-form context, changes in the Excess Bond Premium (EBP) have indeed a far stronger predictive power for economic downturns than for expansions, as noted by Stein (2014) among others. Yet, once we isolate variations in spreads that are caused by monetary policy, we find no evidence of a non-linear transmission mechanism. The bond market response to a monetary shock is economically important, and consistent with Gertler and Karadi (2015), but linear. The same is true of industrial production, employment and unemployment. In short, neither bond investors nor firms appear to react asymmetrically to the decisions of the Fed depending on whether these are expansionary or contractionary. To reconcile the evidence on non-linearity emerging from predictive regressions with the essentially linear nature of the transmission mechanism of monetary policy *via* risk premia we investigate whether other mechanisms might be at play. We find that spreads do indeed respond asymmetrically to macroeconomic surprises that also precede economic downturns. The asymmetry is extreme, in that bad news (i.e. unemployment figures above the average market expectation on the eve of the data release) lead to significant market responses while good news are essentially ignored. The analysis suggests that the reduced-form asymmetry is largely the product of reverse causation. Credit shocks do matter, but most of the abnormal spread fluctuations typically observed ahead of an economic slowdown reflect a predictive rather than a causal relationship. Our results contribute to the literature on the interaction between monetary policy and financial stability (Smets, 2014); corroborate the evidence on the reach-for-yield effects of monetary policy in credit markets (Bekaert et al., 2013; Gertler and Karadi, 2015); and inform the debate on whether these effects could render the exit from a long period of loose money particularly costly (Stein, 2014; Lopez-Salido et al., 2016). Furthermore, our findings are consistent with the idea that deteriorating financial conditions predict rising tail risks for

the economy, or “vulnerable growth”, as in Adrian et al. (2016).

The remainder of the paper is organized as follows. In Section 2 we discuss the literature and use a stylized two-period model to illustrate why an asymmetric transmission mechanism creates a trade-off for monetary policy. In Section 3 we describe the data. In Section 4 we examine a set of forecasting regressions where bond spreads are used as a (potentially non-linear) predictor for various measures of economic activity. We then move to the structural analysis. In Section 5 we sketch our application of the local projection method and our identification strategy for monetary policy shocks. Section 6 presents the associated impulse-responses. In Section 7 we discuss a number of extensions to our baseline structural multivariate model. In Section 8 we study the impact of macroeconomic news on bond spreads. Section 9 concludes.

2 The Dark Side Argument

2.1 Literature

The “dark side” argument describes a causal chain that goes from monetary policy to market risk premia and from these to aggregate economic activity. As Stein (2014) points out, it is the emergence of a non-linearity in this chain that gives rise to a policy trade-off.⁴ The first link in the chain has been studied in the context of the risk taking channel of monetary policy (Borio and Zhu, 2012). Most of the literature takes a banking perspective, exploiting micro data to study the relation between monetary policy and risk taking by financial intermediaries, and finds that banks typically soften their lending standards, demand lower premia and/or engage in riskier investments in periods of easy monetary policy (see Jimenez et al., 2014, and references therein). Analogous mechanisms have been recently found to be active in equity and bond markets (Bekaert et al., 2013; Gertler and Karadi, 2015). Using high frequency data, Gertler and Karadi (2015) find that monetary interventions have a small impact on short-term risk-free rates but a fairly large impact on term and credit risk premia on corporate bonds, and that this second channel accounts for most of their overall macroeconomic effect. Our work is based on similar data – including in particular the Excess Bond Premium of Gilchrist and Zakrajsek (2012) – and on a similar identification of monetary shocks, but focuses on the potentially non-linear nature of the transmission mechanism. Woodford (2012) demonstrates that if risk premia follow non-linear dynamics policy makers indeed face a mean-variance trade-off even in a world where the link between spreads and output is by itself linear, as suggested by Stein (2014). Extending Gertler and Karadi (2015) to a non-linear setup is thus a necessary

⁴“For there to be such a dark side, there would have to be some sort of asymmetry in the unwinding of the effects of monetary policy on these risk premiums, whereby the eventual reversal either happens more abruptly, or causes larger economic effects, than the initial compression” (p.10).

step in our analysis of the dark side argument.⁵

The nexus between financial markets and real economy – the second link in the chain – is the subject of a growing theoretical and empirical literature. There is little doubt by now that in general financial shocks play an important role in causing business cycle fluctuations (Christiano et al., 2014; Nolan and Thoenissen, 2009; Gilchrist and Zakrajsek, 2012; Jermann and Quadrini, 2012; Liu et al., 2013; Gambetti and Musso, 2017). The idea that this linkage might be non-linear, and that changes in credit conditions may have different implications depending on the state of the economy, has clearly gained attention and credibility after the 2008 financial crisis. Mendoza (2010), He and Krishnamurthy (2013), Brunnermeier and Sannikov (2014) develop macroeconomic models where financial shocks are amplified in periods of financial distress, when agents are credit-constrained and thus essentially prevented from fully smoothing consumption. Empirical support for this mechanism is provided by McCallum (1991), Balke (2000) and more recently by Guerrieri and Iacoviello (2013), Alessandri and Mumtaz (2017) and Hubrich and Tetlow (2015), which show that the transmission of various macroeconomic and financial shocks is amplified when financial markets are in turmoil and the economy is close to its borrowing limit. This provides one possible justification for the asymmetry discussed by Stein (2014): increases in bond spreads may have a larger impact on economic activity because they push firms closer to their borrowing constraints. The example developed in Section 2.2 shows in what way an asymmetry of this type can create a trade-off for monetary authorities and change their optimal response to a generic business cycle shock. In essence, the reason is that in a non-linear world a temporary compression in credit spreads has two distinct effects on the future distribution of output: it raises its expected value (for the usual reasons) but it also increases its variance, because the reversal of the spread towards its equilibrium level will cause an even larger output drop at some unknown point in the future. In this situation policy makers may well decide to be relatively more passive and accept a lower expected output level for the sake of (keeping the spreads at their equilibrium level and) reducing volatility.

When dealing with the interaction between asset prices and economic activity one must think carefully about causality. As Stein (2014) acknowledges, a reduced-form asymmetry in the correlation between credit spreads and economic activity could also arise if investors were more sensitive to negative news on the macroeconomic outlook. The asymmetric nature of debt contracts, where bad outcomes are more likely to affect creditors' payoffs than positive outcomes, is a first natural reason for such an asymmetry to arise. Preferences and informational frictions can amplify this asymmetry. Veronesi (1999) presents a model of investment under uncertainty where equity prices systematically overreact to bad news in good times and underreact to good

⁵Chabot (2014) examines the reduced-form dynamics of a range of credit spreads in the US, including EBP, but does not attempt a structural identification of monetary policy or other shocks.

news in bad times. In Epstein and Schneider (2008), ambiguity-averse investors faced with news of uncertain quality behave according to the worst-case scenario, which causes them to react more strongly to bad news than to good news. These asymmetries are documented empirically by Beber and Brandt (2010), who find that contrarian news have a stronger impact on US government bond yields, and that these move the most when bad news occur in good times. Importantly, this type of non-linearity does not cause any complications for monetary authorities. Hence, disentangling it from the alternative credit-constraint mechanism discussed above is critical from both a positive and a normative perspective.

2.2 Mean-variance trade-off in a two-period economy

To see why an asymmetric link between the credit market and the real economy could change monetary policy choices, consider an economy that lasts two periods ($t = 1, 2$) and that is fully characterized by two equations describing respectively the output gap y and the credit spread s :⁶

$$\begin{aligned} y_t &= \gamma \Delta s_t + \xi \Delta s_t I_{\Delta s_t > 0} + e_t \\ s_t &= \rho s_{t-1} + i_t \end{aligned}$$

The output gap is affected by a random disturbance e_t and by the variation in credit spreads relative to the previous period. The impact of rising spreads on the output gap is negative ($\gamma < 0$) and potentially non-linear ($\xi \leq 0$): the $\xi < 0$ case introduces the main asymmetry studied in this paper (though in the empirical analysis we also consider the possibility of a non-linearity in the spread equation itself). The spread follows a simple autoregressive process with persistence $\rho > 0$, and it is affected by the monetary policy rate chosen by the central bank, i_t .⁷ This provides the simplest possible set up where (i) monetary policy works through credit markets, as in Gertler and Karadi (2014); (ii) its effects are temporary; and (iii) the central bank may have to take into account that the economy adjusts non-linearly to a tightening in credit conditions. The set up incorporates a number of extreme assumptions (here monetary policy *only* works through credit markets, the pass-through from the policy rate to the spread is complete, and spreads are not hit by any other shock). These make the trade-off particularly transparent and have no substantive implications for the analysis, which is of course purely qualitative. A monetary stimulus $i_t < 0$ can close the output gap today but it also sows the seeds for the occurrence of a negative gap tomorrow, when the spread reverts back towards its

⁶The example is clearly purely illustrative. Details on the derivations can be found in the Appendix to the paper.

⁷We assume without loss of generality that the equilibrium level of the spread is zero. s_t can equivalently be interpreted as a (zero-mean) 'excess premium' relative to some arbitrary equilibrium level, in line with the EBP used later in our empirical analysis. We also assume throughout $\beta(1 - \rho)(1 + \xi/\gamma) < 1$ to insure that the optimization problem discussed below is well-behaved.

equilibrium level. Consider an economy that starts off from an equilibrium where $y_0 = s_0 = 0$. At time 1 an exogenous shock e_1 takes place, the central bank (CB) observes it and decides whether and how to accommodate it by manipulating i_1 . No actions and no further shocks take place at time 2. Conditional on the shock e_1 , the output gaps at $t = 1$ and 2 are a known function of the policy response:

$$\begin{aligned} y_1 &= \gamma i_1 + \xi i_1 I_{i_1 > 0} + e_1 \\ y_2 &= -(1 - \rho) i_1 (\gamma + \xi I_{i_1 < 0}) \end{aligned}$$

In this world the policy instrument always moves in the opposite direction of the shock, so that the CB chooses to loosen (tighten) if and only if the initial shock is negative (positive). We can thus focus on the case of a recession $e_1 < 0$ that creates an incentive for the CB to implement some monetary stimulus, and study how the optimal size of such a stimulus is affected by risk preferences and non-linearity. The loss function of a **risk-neutral central bank (RN)** is simply the average output gap over the two periods, that is $\ell^{RN}(e, i) = y_1 + \beta y_2$ where β is the CB's discount factor. By replacing y_1 and y_2 and setting $\ell^{RN}(e, i) = 0$, we obtain the optimal risk-neutral choice:

$$i^{RN} = -\frac{1}{\gamma} \left[\frac{1}{1 - \beta(1 - \rho) \left(1 + \frac{\xi}{\gamma}\right)} \right] e \equiv -\frac{\kappa^{RN}(\xi)}{\gamma} e$$

Since $\kappa^{RN} > 0$ and $\gamma < 0$, the policy response has the same sign as the shock and interest rates always fall after a recessionary shock. A myopic or impatient CB fully accommodates the shock: $\beta = 0$ implies $\kappa(\xi) = 1$ and $i = -e/\gamma \equiv i^{FA}$ (where FA stands for full accommodation). This CB chooses to keep the t_1 output gap constant at zero: the future gap will be negative, but it does not care about it. Full accommodation is also optimal if the spread is a random walk, as $\rho = 1$ again implies $\kappa = 1$. If there is no mean-reversion, the shock can be fully neutralized without paying any costs at $t = 2$. More generally, however, the CB *overreacts* to the shock:

Result (1) A risk-neutral CB responds aggressively to the shock: ($\beta \neq 0$, $\rho < 1$) imply $\kappa^{RN}(\xi) > 1$ and thus $i^{RN} < i^{FA}$. Furthermore, the policy response is increasing in the absolute magnitude of the non-linearity, i.e. decreasing in ξ : $\partial \kappa^{RN}(\xi) / \partial \xi < 0$.

Knowing that the stimulus comes at the cost of a future contraction, a risk-neutral CB simply engineers a positive gap today that exactly compensates for the (discounted) negative gap that will materialize tomorrow. The existence of a non-linearity does not change the nature of this problem: it simply makes the CB more aggressive. This behavior creates significant volatility in y – effectively a boom followed by a recession – but by construction the CB is not concerned about it. A **risk-averse central bank (RA)** aims instead to minimize the

variance of the output gap around its zero target. The loss function is given in this case by $\ell^{RA}(e, i) = y_1^2 + \beta y_2^2$. Setting $\partial \ell^{RA}(e, i) / \partial i = 0$ gives the following unique solution:

$$i^{RA} = -\frac{1}{\gamma} \left[\frac{1}{1 + \beta(1 - \rho)^2 \left(1 + \frac{\xi}{\gamma}\right)^2} \right] e \equiv -\frac{\kappa^{RA}(\xi)}{\gamma} e$$

For a myopic central bank, or one that faces random-walk spreads, $\kappa^{RA}(\xi) = 1 = \kappa^{RN}(\xi)$, so the solution is again full accommodation, $i^{FA} = -e/\gamma$. In this case, however, if we move away from those extremes we find that the CB accommodates the shock only in part:

Result (2) A risk-averse CB responds mildly to the shock: $(\beta \neq 0, \rho < 1)$ imply $\kappa^{RA}(\xi) < 1$ and thus $i^{RA} > i^{FA}$. Furthermore, the policy response is decreasing in the absolute magnitude of the non-linearity, i.e. increasing in ξ : $\partial \kappa^{RA}(\xi) / \partial \xi > 0$.

Note first that $\kappa^{RA}(0) = (1 + \beta(1 - \rho)^2)^{-1} < 1$. Even in a linear world ($\xi = 0$) mean-reverting credit spreads create a cost in terms of volatility that a risk-averse CB naturally takes into account when taking its decision. The mean-variance trade-off is such that, in general, the CB accepts a negative average gap for the sake of keeping volatility under control. More importantly, the shape of the trade-off is a function of the non-linearity. The larger is ξ in absolute terms (i.e. the lower is $\xi < 0$), the larger is the cost in terms of variance that must be paid to stabilize today's output gap, and the lower is $\kappa^{RA}(\xi)$.

This example allows us to draw two main conclusions. First, mean reversion in credit spreads creates by itself a mean-variance trade-off that makes a risk-averse central bank more cautious in tackling negative economic shocks. A full accommodation of the shock is generally suboptimal for a risk-averse authority. Second, the terms of the trade-off, and the optimal degree of accommodation, depend on the structure of the economy. The central bank's incentive to counter recessionary shocks is weaker if a reversal in spreads has a stronger impact on the economy than their initial fall. This provides an intuitive formalization of the argument by Stein (2014).

3 Data

We study credit and output dynamics in the United States between 1973 and 2012. Our main proxies of credit conditions are the corporate bond spread and the Excess Bond Premium (EBP) provided by Gilchrist and Zakrajsek (2012) (henceforth GZ). Using bond-level data, GZ construct measures of the spreads paid by US corporations over risk-free rates of the corresponding maturities. These are subsequently split through a regression model into a predicted component, driven by firm-specific and macroeconomic factors affecting the expected

default probability of the issuer, and an additional component (EBP) that measures residually the excess return required by investors over and above the compensation for credit risk. A simple cross-sectional average of these two variables then provides economy-wide measures of expected credit spread and EBP.

From our perspective these indicators have two important advantages. First, they are theoretically appealing, as they do not suffer from the maturity mismatch that plagues commonly used measures such as the difference between yields on BAA bonds and some benchmark risk-free rate. This separation between term premia and credit risk premia is particularly important in our case because the risk-taking implications of monetary policy involve the latter but not the former. Second, GZ demonstrate that the spread – and particularly the EBP component – has significant predictive power for economic activity. This allows us to test for non-linearities in a set-up where there is sound evidence of a significant baseline (linear) correlation between the variables of interest. The GZ spreads are displayed in Figure 1, together with the BAA-over-AAA bond spread calculated by Moody’s, for reference. To measure economic activity we rely on three standard indicators, namely the industrial production index, non farm payroll employment and the unemployment rate. We also use in all our specifications the one-year government bond rate, US CPI and the term spread defined as the difference between ten-year and three-month constant-maturity Treasury yield.⁸ In the discussion of our findings we also use the Chicago Financial Conditions index as alternative to the EBP, the Michigan Index of consumer confidence and the uncertainty measure constructed by Jurado et al. (2015) (see Section 7). All data are from FRED St. Louis, apart from the latter two, downloaded by the site of the University of Michigan and the website of Serena Ng, respectively.

The last important piece of our information set is a measure of macroeconomic news constructed with daily data, taken from Bloomberg. For every month in the sample, we calculate an “unemployment news” indicator as the difference between the unemployment rate update published in that month and the corresponding market expectation on the day before the data release. The latter is obtained as the median across Bloomberg forecasters. In Section 8 we use this indicator to look at the dark side argument from a reverse angle, asking how news shocks that might lead bad turns in the business cycle affect bond markets.

4 Predictive regressions

We start with a thorough check of the key stylized fact that motivates our research question – namely the existence of an asymmetric reduced-form relation between bond spreads and economic activity. To investigate it, we use predictive regressions of the following type:

$$\nabla^h Y_{t+h} = a(L)\Delta Y_t + \beta_2 term_t + \beta_3 realr_t + \beta_3 \hat{S}_t^{GZ} + \beta_4 EBP_t + \beta_5 EBP_t^+ + \epsilon_{t+h}, \quad (1)$$

⁸These variables are necessary in order to mimic the baseline linear framework by GZ.

where the dependent variable $\nabla^h Y_{t+h}$ is the (cumulative) percentage change in economic activity between t and $t+h$, $a(L)\Delta Y_t$ is a distributed lag of the dependent variable, $term_t$ is the term spread, defined as the difference between ten-year and three-month constant-maturity Treasury yield, and $realr_t$ is the short term real interest rate. The term $S_t^{\hat{G}Z}$ is the predicted corporate bond spread while EBP_t is the Excess Bond Premium, both taken from GZ (see Section 3 for details). This regression model essentially augments the set-up studied in GZ with a *local peak* transformation of the excess premium, EBP_t^+ , that allows EBP to have a potentially asymmetric "effect" depending on the direction in which the spreads are moving. For a generic variable x_t , the local peak function is defined as follows:

$$x_t^+(h) = x_t I[x_t > \max(x_{t-1}, x_{t-2}, x_{t-3}, \dots, x_{t-h})] \quad (2)$$

where $x_t^+(h)$ equals 0 if x_t is not a peak over the past h periods, x_t otherwise. By introducing EBP_t^+ in the regressions we can capture shifts in the correlation between spreads and output that take place when EBP_t reaches a local maximum. A maximum can be reached either because the spread rises consistently for h periods or because a large shock suddenly pushes it above its recent historical values. The h parameter determines how persistent the shock to the EBP has to be for this additional regressor to be activated. Note that with $h = 1$ any increase in EBP qualifies as a local peak, giving the special case of a 'pure' sign asymmetry considered in Stein (2014). As h increases, non-zero values of EBP_t^+ become progressively less frequent, capturing only large/persistent movements in spreads. The economic rationale for using this transformation is that small, temporary shocks to credit conditions can be more easily smoothed out by firms through profit margins, while large or persistent changes in the cost of credit are more likely to affect investment and output.⁹ In this setup, one can test for asymmetric effects by testing whether β_5 is significantly different from zero. This approach to testing for asymmetries has been used extensively to study the effects of oil price shocks on economic activity and inflation (Borenstein et al., 1997; Meyler, 2009). Its methodological limitations are discussed by Kilian and Vigfusson (2013).¹⁰

The results obtained from these predictive models are reported in Table 1. The table collects a range of specifications that differ along three dimensions: (i) the lags used to compute local peaks in EBP (from 12 to 36); (ii) the forecasting horizon (from 6 to 18 months ahead); and (iii) the measure of economic activity used as forecasting target (employment, industrial production, unemployment rate). EPB confirms to be a significant predictor of economic ac-

⁹The nature of EBP^+ is discussed further in Section 7, where we also consider alternative ways of capturing the non-linearity of interest.

¹⁰Kilian and Vigfusson (2013) show that, in the case of censored variables, the regression coefficients can give a very distorted view of whether or not the shocks are transmitted asymmetrically, and that the bias can go either way (with small coefficients masking significant dynamic asymmetries, or *vice versa*). The structural multivariate models introduced in Section 5 are less prone to this problem.

tivity, as in GZ. This result is robust across specifications and measures of economic activity. The coefficients display the expected signs: an increase in real interest rates, a rise in credit spreads (either in its predicted component or in EBP) and a flattening of the yield curve are all associated to a future economic contraction. The key object of interest is of course EBP_t^+ . In the case of industrial production, the coefficient is highly significant for all horizons and local peak specifications. In the employment and unemployment regressions the coefficient is again significant as long as one focuses on large values of h – i.e. persistent increases in credit spreads – and forecasting horizons of at least 12 months. These regressions results support and extend the evidence presented in Stein (2014), where a similar predictive exercise is carried out using a different sample, using a simple(r) dummy variable specification and GDP growth as a target. All in all, the evidence clearly supports the notion that credit spreads move more ahead of a slowdown. In the remainder of the paper we turn to the question of what are the structural causes of this non–linearity.

5 A non–linear model of the monetary transmission mechanism

To examine the causal chain underpinning the dark side argument one needs to move from univariate reduced-form regressions to a model that (i) captures feedbacks between real economy, financial markets and monetary policy, (ii) can be combined with a credible structural identification strategy, and (iii) allows for a non–linear propagation of monetary shocks. To that end, we resort to the following system of equations:

$$y_{t+h} = a^h + \sum_{i=0}^{p-1} B_i^h y_{t-i} + \Theta^h L y_t^+ + v_{t+h} \quad (3)$$

This system models a set of endogenous variables y_t as a function of their own lags plus their local peak transformations y_t^+ . The matrix L is a selection matrix that contains only zeros except for a single unit entry corresponding to the variable and equation for which the asymmetry is assumed to be relevant.¹¹ The *local projections* method by Jordà (2005) can be promptly used to calculate generic impulse-response functions in this context. In particular, given a shock of interest, the IRFs can be computed taking into account the local peak functions as $IRF(h, t) = B_0^h d_i + \Theta^h L \tilde{y}_t$, where $\tilde{y}_t = [(y_t + d_i)^+ - y_t^+]$ and d_i is a shock to the i^{th} variable in the system. Notice that in the first term in \tilde{y}_t the local peak function is applied to the sum of y_t and the shocks vector d_i : this makes the IRF dependent on the *history* of the variable and on

¹¹For example, in a bivariate model allowing for asymmetric effects of the second variable in the first equation we have $L = \begin{pmatrix} 0 & 1 \\ 0 & 0 \end{pmatrix}$, so that $L y_t^+ = \begin{pmatrix} y_{2,t}^+ \\ 0 \end{pmatrix}$. The specification of L is discussed in detail below.

both the *sign* and *size* of the shock, highlighting the complex non-linearities that might arise in systems that include censored variables. In our baseline specification the y_t vector includes economic activity (measured again by industrial production, employment or the unemployment rate), EBP, term spread, (log) CPI index and the interest rate on one-year government bonds. We later extend the analysis to include alternative credit indicators and measures of volatility and sentiment.

By altering the definition of L we can examine two separate ways in which credit markets might render the transmission of monetary policy asymmetric. In the first case we assume the spread to respond linearly to monetary surprises, but allow economic activity to respond asymmetrically to variations in credit spreads. Here the L matrix has an entry equal to 1 in the position that corresponds to the spread term in the output equation. In the second case we allow for a non-linearity arising within financial markets, and let the spread respond asymmetrically to monetary shocks. Here the L matrix is set up with an entry equal to 1 in the position that corresponds to the monetary policy indicator in the spread equation.¹² From an economic standpoint, these two cases represent situations where credit constraints affect two different sets of agents: nonfinancial firms and bond investors. In particular, an asymmetric output response would be consistent with credit-constrained firms being relatively more sensitive to an increase in their funding costs (measured by EBP) than to a decrease in such costs. An asymmetric EBP response could instead arise if the constraints affected bond investors, which would be forced to dump the bonds after an increase in their own funding costs (proxied by the monetary policy rate), causing EBP to be relatively more sensitive to monetary contractions. This represents a scenario where the “risk off” phase triggered by a contractionary monetary surprise is more dramatic or more abrupt than the “risk on” phase, or, put differently, investors buy risk gradually but offload it quickly when monetary conditions tighten. Like the previous one, this asymmetry stems from the interaction of leverage and occasionally binding borrowing constraints (Brunnermeier and Sannikov, 2014). Its empirical relevance has already been documented for currency markets, where the conventional wisdom that ‘exchange rates go up by the stairs and down in the elevator’ is supported by formal econometric evidence (Brunnermeier et al., 2009). Our set up allows us to test its relevance in the case of corporate bond markets.

To identify monetary policy shocks we exploit high-frequency financial data as external instruments, following Gertler and Karadi (2015) (see also Stock and Watson, 2012; Mertens and Ravn, 2013).¹³ The methodology consists of four steps. First, we estimate the reduced form one

¹²In principle these two non-linearities could be combined in a single model. However, testing them one at a time is both convenient and more informative: introducing them together would complicate the estimation and the interpretation of the results and weaken the inference.

¹³Caldara and Herbst (2016) highlight the importance of exploiting this identification when looking at the relevance of monetary policy shocks for economic activity using a (linear) Bayesian Proxy SVAR model.

step ahead residuals using OLS regressions on the multivariate model. Second, we instrument the residual of the interest rate equation (for which we use the one-year government bond rate¹⁴) with changes in asset prices that occur on the days of the FOMC meetings. The presumption is that these are correlated with the unexpected component of the observed variation in policy rates but uncorrelated with other confounding factors.¹⁵ Third, we estimate the contemporaneous response of the remaining variables to the fitted values obtained in the first stage. This identification strategy is attractive in our context because it copes well with simultaneity problems (monetary conditions are likely to respond to financial developments within a given month) and with the fast response of financial markets to central banks' decisions (an issue that is likely to be particularly important for the unconventional monetary interventions undertaken by the Fed after the Global Financial Crisis). Fourth, having estimated the *contemporaneous* impact of monetary policy shocks on all the variables in the system, we use local projections to measure the *dynamic* impact of the shock taking into account possible non-linearities in its transmission mechanism. In the robustness analysis of Section 7 we also consider recursive identification schemes.

6 Bond markets and monetary policy shocks

The first proposition we examine is whether an increase in spreads triggered by an unexpected monetary tightening causes a disproportionately large response in economic activity. We consider a parsimonious specification of the model that includes economic activity (variously defined, see below), CPI inflation, the one-year government bond rate, the term spread and the EBP. The local peaks in EBP are defined as in Section 4 and allow for a different transmission of the shock when the spread is rising.¹⁶ The responses generated by the model are displayed in Figure 2. The plots are organized as follows. Each row refers to a specification that includes a different measure of economic activity. Moving down from the top, these are employment, industrial production and the unemployment rate. The columns display the responses associated with shocks of different magnitudes: moving from left to right, these rise from 25 to 100 basis points. Within each plot, the black line represents the estimated median responses to a restrictive monetary shock (i.e. an unexpected increase in the one-year government bond rate) and the red line represents the response to an expansionary shock of the same size. The latter are multiplied by minus one to facilitate the visual comparison of the effects

¹⁴We choose the one-year maturity because it strikes a good balance between (i) being sufficiently sensitive to monetary surprises (so that the instruments used in the first stage are valid) and (ii) accounting for term structure (i.e. forward guidance) effects. See Gertler and Karadi (2015) for details.

¹⁵Candidate instruments are 1 month fed fund future rates, 3 month fed fund future rates and Euro/Dollar deposits 2, 3, and 4 months ahead. However, we use only the 3-months ahead fed funds rate in the estimation, as it is the most powerful in the first stage.

¹⁶We calculate the peaks over an horizon of $h = 12$ months. The choice of h is however not important: our findings are robust to different values – see Section 7.

of positive and negative shocks. In all cases the median estimated response is accompanied by 68% and 90% confidence intervals, displayed respectively as dark and light grey areas. A 25 basis point increase in the policy rate generates a contraction in employment and industrial production and an increase in the unemployment rate. These effects are persistent and highly significant. Crucially, the IRFs obtained from a positive and a negative shock overlap almost perfectly: there is no evidence that contractionary shocks have a larger impact on economic activity. Although one could suspect that this result depends on the modest size of the shock, this turns out not to be the case. If we condition on shocks of 50 basis points (column 2) or or 100 basis points (column 3), the estimated responses naturally become larger but the equivalence between positive and (inverted) negative shocks is confirmed for all forecasting horizons.

There is a second possibility for a policy trade-off to arise: contractionary monetary policy shocks may force bond investors to reduce their leverage, implying that EBP moves relatively more in response to unexpected increases in the one-year T-Bill rate (see Section 5). This would clearly make the transmission asymmetric even in a world where the linkage between bond spreads and economic activity in itself is completely linear. To investigate this possibility, in Figure 3 we show the response of EBP to a monetary policy shock, again identified via external instruments, from a model where the monetary policy indicator enters non-linearly the EBP equation. As in the previous case, the three columns report the responses associated with shocks of different sizes and the rows refer to specifications based on our three alternative measures of economic activity. The dynamics in EBP are very similar across specifications. In particular, the behavior of the spread is consistent with that documented by Gertler and Karadi (2015): EBP increases significantly on impact and remains positive for over a year after the shock. This result adds to the existing evidence in support of a market-based risk-taking channel for monetary policy in the US. Our model also allows us to make statements concerning the way markets adjust to shocks of different size and direction. In short, neither of the two dimensions matters. Moving across the columns of figure 3, one clearly sees that the response by EBP is linear in the size of the monetary shock irrespective of which economic activity indicator is used in the model. Furthermore, the overlap between the red and gray bands is striking: whatever the size of the shock, the responses of EBP to contractionary and expansionary monetary surprises are nearly identical. Given this result we do not investigate this alternative channel any further.

7 Discussion

Below we extend our baseline structural analysis along various dimensions, considering in turn: the role of economic uncertainty and confidence, the nature of the transmission during recessions, different indicators of financial conditions, alternative non-linear transformations

of the EBP, and a broader class of "credit shocks" identified by a simple recursive ordering assumption. We summarize these alternative specifications in Table 2. For brevity we only report some of the results obtained from these analyses; the remaining figures are available in the Supplementary Appendix B.

7.1 Accounting for uncertainty and consumer confidence

In the aftermath of the Global Financial crisis there has been a growing interest in evaluating the effects of uncertainty on the business cycle (see Bloom, 2014, for a survey). The literature shows that the impact of uncertainty depends on the state of the real economy (Caggiano et al., 2014) and might interact in a non-trivial way with financial market dynamics as documented by Caldara et al. (2016). Although we are not interested in identifying the effects of uncertainty *per se*, the interaction between uncertainty and credit conditions might play a role in our case. Hence, we include among the controls the measure of aggregate uncertainty recently proposed by Jurado et al. (2015) and the Michigan Index of consumer confidence, a widely-used series in the literature on expectations, news and confidence shocks (Barsky and Sims, 2011, 2012). The forward-looking nature of this indicator can be useful in multivariate models also because it broadens the information set included in the model. Despite a widening of the gap between the underlying medians, the confidence bands associated with expansionary and contractionary monetary policy shocks do overlap, thus validating the conclusion that the transmission mechanism is linear, as shown in Figure 4.

7.2 Recessions, broader financial conditions indicators, alternative EBP transformations

A growing body of research evaluates the effects of monetary policy in different phases of the business cycle. Santoro et al. (2014) find that monetary policy exerts stronger effects on output during contractions, while Tenreyro and Thwaites (2016) find that contractionary policy shocks have larger effects than expansionary shocks. Although we do not develop a formal state-dependent model, our framework allows us to condition the estimate of our IRFs over periods of recessions instead of taking into account the whole history of economic activity (the IRFs indeed depend not only on size and sign of the shock but also on the history of the system, see Section 5). We thus replicate the analysis illustrated in Figure 2 conditioning on periods when the economy is in recession instead of conditioning simply on the mean value of the endogenous variables as in the baseline case. In particular, we average over "recessions", defined as periods in which the year-on-year change in employment, industrial production or unemployment is negative (positive for the latter). Again, we do not find any evidence of asymmetry.

Our choice to focus on EBP reflects its proven predictive power (Gilchrist and Zakrajsek, 2012) and its centrality in the argument developed in Stein (2014). Other financial variables may however influence the business cycle, possibly in a non-linear fashion (Adrian et al., 2015; Hubrich and Tetlow, 2015). To evaluate this possibility, we replace EBP with the Chicago Fed Financial Condition Index, a model-based indicator that provides a broader picture on funding conditions in money, debt and equity markets. Again, we find that the effects of monetary easing and tightening are not statistically distinguishable from one other. This leads us to exclude the possibility that other markets embed non-linearities that are not observed in the corporate bond segment.

Our analysis relies on the identification of “exceptional” increases in bond spreads, which is inevitably arbitrary to some extent. With respect to the baseline definition of EBP^+ employed thus far, two alternative options can be considered: (i) the horizon over which the net increase is computed and (ii) the type of non-linear function used in the calculation. As for the first dimension, we compute our local peaks over a period of 24 and 36 months instead of 12 as in the baseline case. The results are qualitatively and quantitatively unaffected. Moving to the functional forms, we test two alternatives. The first one, labeled S_{diff} , is defined as $x_t^+(j) = (x_t - x_{t-h})I[x_t - x_{t-h} > 0]$. Here the net increase is activated every time the change in EBP over the last $j = t - h$ periods is positive (where, again, we choose $h=12$ as baseline). The second one, labeled S_{plus} , is defined as $x_t^+ = x_t I[x_t > 0]$: this restricts the focus to occurrence of positive excess bond premia. The indicators are displayed in Figure 5, where we compare our baseline EBP^+ indicator (top panel) to S_{diff} and S_{plus} (middle and bottom panel). The IRFs obtained using these alternative transformations of the EBP show that, although minor differences between positive and negative shocks emerge in the case of S_{plus} , the non-linearity is never statistically significant.

All of the results discussed in this subsection are available in the Supplementary Appendix B.

7.3 Credit shocks

A final concern might be related to the quantitative relevance of the monetary policy shocks on which we build our analysis in Section 6. Monetary shocks are clearly central to the "dark side" argument of Stein (2014). Furthermore, they can be identified with a good degree of confidence through the high-frequency identification proposed by Gertler and Karadi (2015). Yet, one could think that these shocks explain a portion of the variance of EBP in the data that is simply too small for a statistically significant non-linearity to arise. Put differently, there might be other shocks that, being more important drivers of EBP, would stand more chances of revealing a non-linear causal relation between bond markets and the real economy. To take this possibility into account, we go back to the model of Section 5 but shift the focus

from monetary shocks to a more encompassing class of “credit shocks”. These are identified recursively by assuming only that economic activity responds with a lag to EBP. Hence, we take as exogenous all variations in EBP that are orthogonal to the state of the business cycle at time t . The effect of these shocks is plotted in Figure 6. In this case black and red lines represent respectively the median responses to a rise and a fall in EBP. The structure of the figure is otherwise identical to that of Figure 2.¹⁷ Two considerations are in order. First, the peak responses are somewhat larger than those in Figure 2 (the contemporaneous responses are zero by construction). Second, although the overlap between responses is again perfect for small or medium changes in EBP (columns 1 and 2), some asymmetries do show up in the case of large 100bps shocks (column 3). In this case output appears to respond more when EBP rises. These results can be interpreted in two ways. One is that some of the primitive shocks captured in this exercise – variations in risk preferences being an interesting candidate – are *both* more powerful than monetary policy shocks in shifting EBP *and* more likely to induce an asymmetric output response. Another one is that the Cholesky decomposition ends up mixing proper shocks with endogenous market responses to unobserved events that affect EBP (simultaneously) and economic activity (with some delay). Given the evidence that we discuss in the next Section, we favor the latter.

8 Bond markets and macroeconomic news

Our analysis suggests that, although credit markets predict recessions more accurately than expansions (Section 4), they absorb and transmit monetary policy shocks in a perfectly linear way (Sections 6 and 7). The reduced-form asymmetry must thus have a different explanation. An intuitive candidate comes from asset pricing: due to their payoffs, preferences and/or information sets, investors may be more sensitive to negative news on the macroeconomic outlook, causing spreads to move more ahead of a recession (see Section 2.1). A closer look at this statement reveals that three conditions are needed for this asymmetry to explain our results. First, there must be macroeconomic “news shocks” that – controlling for a number of covariates – have significant predictive power for future economic activity. Second, these shocks must have an asymmetric impact on credit markets. Third, they must be orthogonal to monetary policy shocks, which implies that they are not detected by the structural analysis presented in Section 6. To check whether this explanation holds in our data, we resort to our proxy of “news” regarding the US unemployment rate, U_t^{news} . This is defined as the difference between the unemployment rate update published in month t and the median unemployment

¹⁷Note that the shocks are defined in terms of basis points, rather than standard deviations, to allow for a more direct comparison across specifications.

expectation held by market participants the day before the data release (see Section 3).¹⁸

In order to check whether unemployment news contain useful information for market participants – i.e. whether they forecast future economic activity over and above a standard set of macroeconomics controls (the first of the conditions mentioned above) – we estimate a set of predictive regressions analogous to those used in Section 4. More specifically, we estimate models of the following type:

$$\nabla^h Y_{t+h} = a(L)\Delta Y_t + \Gamma' x_t + \beta U_t^{news} + \epsilon_{t+h} \quad (4)$$

where Y is a measure of economic activity (unemployment rate, industrial production or employment), x_t is a vector of controls that includes term spread, real fund rates and the predicted corporate default spread (GZ spread), and U_t^{news} is the unemployment rate news indicator. The news coefficient β has the expected sign in all specifications: “bad news” predict higher unemployment, lower industrial production, and lower employment 6, 12 and 18 months ahead. In Table 3 we report the robust t -statistics associated to the coefficient: this is above 1.66 in absolute terms (its 90% critical value) in more than half of the specifications, indicating that the marginal predictive power of U_t^{news} is statistically significant. The second question is whether credit spreads react more to bad than to good news. We test this proposition using the following equation:

$$EBP_t = \alpha + \phi(L)EBP_t + \beta(L)^{bad} U_t^{news} I(U_t^{news} > 0) + \beta(L)^{good} U_t^{news} I(U_t^{news} \leq 0) + \Gamma' x_t + u_t \quad (5)$$

where x_t is the usual vector of controls (now including the lagged unemployment rate as well as term spread, real fed fund rates and predicted GZ spread) and $\phi(L)$ and $\beta(L)$ represent lagged polynomials. These capture the persistence of EBP and the occurrence of possible misalignments between the Bloomberg data releases and the credit risk priced by the EBP.¹⁹ Indeed, Bloomberg releases on surprises occur at the beginning of the month, while data underlying the estimation of the EBP by Gilchrist and Zakrajsek (2012) are month-end secondary market prices. Hence, if interested in pricing the credit risk p_{jt} of the j -th bond at month t , exploiting information coming from Bloomberg releases beyond the possibly relevant data, one will only be able to rely on $news_{t-1}$ in his/her information set. To separate the effects of good and bad news we use a simple dummy variable that separates the months when the news are bad ($U_t^{news} > 0$, meaning a higher-than-expected unemployment rate) from those when they are good (the opposite). The results are displayed in Table 4. The first two columns show estimated coefficients and p-values from a specification without the z_t controls, while the last two

¹⁸Macroeconomic surprises are commonly used in the asset pricing empirical literature, see e.g. Gürkaynak et al. (2005), Faust et al. (2007), Goldberg and Grisse (2013).

¹⁹It turns out that two lags are sufficient to accurately model EBP and only one lag of the *news* variable emerges as significant.

include these controls. The signs are again in line with expectations: good news lower EBP while bad news increase it. However, only $\hat{\beta}^{bad}$ is significantly different from zero. EBP rises significantly in response to bad macroeconomic news, but it is overall unresponsive to good news. This is a crucial piece of information. Not only does it confirm that financial markets use a broader information set than that typically included in a VAR; but, more importantly, it also highlights that an econometrician using the VAR might hold the markets accountable for "shocks" that are instead a response to changes he or she does not observe. This is ultimately the difference between investors predicting and causing a recession.

To close our argument we also check that the macro news we are considering are orthogonal to the surprises in asset prices that we use to identify monetary policy shocks. From an economic perspective this condition is perhaps not as important as the previous ones. From a statistical perspective, however, it is interesting for two reasons. First, if the shocks turned out to be correlated one would suspect that they are not shocks after all – or that at least one of them is not. Second, a low correlation would help explain why unemployment news do not trick the structural model of Sections 5 and 6 into wrongly concluding that the monetary transmission mechanism is non-linear. Our last test is then based on the following simple regression

$$U_t^{news} = \alpha + \lambda(L)MP_{surprise} + u_t \quad (6)$$

The results, summarized in Table 5, indicate that macro news and monetary shocks are indeed orthogonal.²⁰

9 Conclusions

Monetary policy exerts a significant influence on credit risk premia, and the correlation between risk premia and future economic activity appears to be asymmetric: spikes in credit spreads are often followed by significant economic contractions, but the linkage appears to be far weaker when the spreads decline. If it interfered with the transmission mechanism, this asymmetry would present monetary authorities with a delicate trade-off, as a monetary stimulus aimed at easing credit conditions and closing the output gap in the short run would also increase the risk of a costly reversal in financial markets in the longer term (Stein, 2014). After sketching a simple analytical example to illustrate the nature of the policy trade-off, we develop a thorough econometric investigation of the non-linearity, its structural causes, and its implications for

²⁰A similar analysis based on surprises on payroll employment, gives less clear-cut results. In particular, the predictive content for economic activity of this measure is less stable and strong. This looks consistent with Goldberg and Grisse (2013), who argue that this variable has a different impact in times of low and high risk, which, in our framework, is somewhat picked-up by the volatility of the EBP, particularly high in the final part of the sample. A more thorough analysis of the properties of various "news" measures should be based on daily or intra-daily data rather than on monthly data and consider the change in asset prices on the exact day of the macroeconomic news. Such extensions are beyond the main objective of this paper.

monetary policy. We use monthly data and rely on the corporate bond spread series provided by Gilchrist and Zakrajsek (2012) to capture credit conditions. We document that, in a reduced-form set-up, the relation between the excess bond premium (EBP) and economic activity is indeed highly non-linear: over the last four decades, EBP has systematically experienced sharp rises ahead of slowdowns in economic activity, consistent with the more recent experience of the Great Recession. However, we find that the asymmetry has little to do with monetary policy. Monetary shocks – which we identify exploiting high-frequency financial data as in Gertler and Karadi (2015) – have a symmetric impact both on EBP and on economic activity. On the other hand, bond markets respond in a strongly asymmetric way to macroeconomic news: the release of unexpectedly high unemployment figures triggers a sharp upward revision in EBP, while positive surprises leave it unaffected. Taken together, the results suggest that the asymmetry observed in the data arises because bond investors are (not surprisingly) more sensitive to bad news that anticipate a deterioration in the macroeconomic outlook. Since this phenomenon does not depend on their decisions, central banks do not need to alter their monetary policy frameworks in any particular way. If applied to the current US outlook, our analysis suggests that the lift-off from the prolonged monetary expansion implemented by the Fed is unlikely to come at a cost that is so high as to raise doubts on whether the stimulus was worth undertaking in the first place.

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Table 1: CREDIT SPREADS AND ECONOMIC ACTIVITY: NON-LINEAR GZ REGRESSIONS

Order of local peak	12			24			36		
Forecast horizon	6	12	18	6	12	18	6	12	18
Employment									
Term Spread	-0.09	-0.19	-0.31	-0.09	-0.21	-0.30	-0.09	-0.19	-0.31
<i>p-val</i>	0.03	0.06	0.01	0.04	0.01	0.01	0.04	0.06	0.02
Real Fed Funds	0.09	0.14	0.14	0.09	0.14	0.14	0.09	0.15	0.14
<i>p-val</i>	<i>0.20</i>	<i>0.37</i>	<i>0.28</i>	<i>0.19</i>	<i>0.21</i>	<i>0.25</i>	<i>0.18</i>	<i>0.30</i>	<i>0.24</i>
Predicted GZ spread	-0.10	-0.13	-0.21	-0.10	-0.13	-0.20	-0.09	-0.12	-0.20
<i>p-val</i>	<i>0.11</i>	<i>0.33</i>	<i>0.21</i>	<i>0.11</i>	<i>0.25</i>	<i>0.18</i>	<i>0.13</i>	<i>0.32</i>	<i>0.20</i>
EBP	-0.14	-0.17	-0.16	-0.15	-0.18	-0.17	-0.14	-0.17	-0.17
<i>p-val</i>	0.00	0.00	0.01	0.00	0.00	0.00	0.00	0.00	0.00
<i>EBP</i> ⁺	-0.13	-0.20	-0.21	-0.14	-0.23	-0.24	-0.16	-0.25	-0.25
<i>p-val</i>	0.13	0.12	0.11	0.16	0.02	0.01	0.15	0.05	0.01
<i>R</i> ²	0.76	0.65	0.56	0.76	0.64	0.56	0.76	0.65	0.56
Industrial production									
Term Spread	-0.06	-0.11	-0.19	-0.06	-0.11	-0.19	-0.06	-0.11	-0.19
<i>p-val</i>	<i>0.22</i>	<i>0.29</i>	<i>0.17</i>	<i>0.26</i>	<i>0.39</i>	<i>0.14</i>	<i>0.24</i>	<i>0.39</i>	<i>0.11</i>
Real Fed Funds	0.01	0.00	-0.06	0.02	0.01	-0.06	0.02	0.01	-0.06
<i>p-val</i>	0.89	0.97	0.58	0.83	0.93	0.54	0.81	0.92	0.53
Predicted GZ spread	-0.13	-0.16	-0.24	-0.12	-0.16	-0.24	-0.12	-0.16	-0.24
<i>p-val</i>	0.03	<i>0.11</i>	0.04	0.03	<i>0.11</i>	0.02	0.03	0.16	0.02
EBP	-0.20	-0.19	-0.15	-0.21	-0.20	-0.16	-0.21	-0.20	-0.16
<i>p-val</i>	0.01	0.03	0.10	0.00	0.01	0.06	0.00	0.01	0.07
<i>EBP</i> ⁺	-0.22	-0.29	-0.22	-0.26	-0.33	-0.26	-0.26	-0.34	-0.26
<i>p-val</i>	0.01	0.00	0.02	0.01	0.00	0.00	0.01	0.00	0.00
<i>R</i> ²	0.45	0.35	0.26	0.46	0.36	0.26	0.46	0.36	0.26
Unemployment rate									
Term Spread	0.12	0.25	0.36	0.12	0.25	0.35	0.12	0.25	0.35
<i>p-val</i>	0.07	0.03	0.77	0.08	0.03	0.10	0.09	0.03	<i>0.48</i>
Real Fed Funds	-0.06	-0.07	-0.02	-0.07	-0.08	-0.03	-0.07	-0.08	-0.03
<i>p-val</i>	<i>0.63</i>	<i>0.76</i>	<i>0.99</i>	<i>0.62</i>	<i>0.74</i>	<i>0.96</i>	<i>0.61</i>	<i>0.70</i>	<i>0.97</i>
Predicted GZ spread	0.10	0.11	0.17	0.09	0.11	0.17	0.09	0.10	0.16
<i>p-val</i>	<i>0.27</i>	<i>0.48</i>	<i>0.81</i>	<i>0.28</i>	<i>0.47</i>	<i>0.56</i>	<i>0.30</i>	<i>0.48</i>	<i>0.57</i>
EBP	0.26	0.24	0.19	0.25	0.24	0.20	0.25	0.24	0.19
<i>p-val</i>	0.00	0.00	<i>0.80</i>	0.00	0.00	0.01	0.00	0.00	<i>0.25</i>
<i>EBP</i> ⁺	0.09	0.23	0.24	0.13	0.29	0.31	0.16	0.32	0.33
<i>p-val</i>	<i>0.30</i>	0.05	<i>0.86</i>	<i>0.21</i>	0.01	0.04	<i>0.17</i>	0.00	0.00
<i>R</i> ²	0.53	0.44	0.38	0.54	0.44	0.39	0.54	0.45	0.39

Notes: Sample is 1973:01 - 2012:12. Dependent variables is $\nabla^h Y_{t+h}$, where Y_t denotes the respective economic activity variable in the sub panel title in month t and h is the forecasting horizon. *Order of local peak* represents the number of periods over which the asymmetric term of the financial variable is computed. Each regressions also include a constant and p lags of the dependent variable (not reported), where p is chosen by the BIC. Entries in the table denote the standardized estimates of the OLS coefficients associated with each financial indicator, whereas italics terms are the p -values computed by means of the Newey–West (1987) correction. [

Table 2: MULTIVARIATE STRUCTURAL MODEL: LIST OF ALTERNATIVE SPECIFICATIONS

#	Difference with respect to the baseline model
1	Includes Confidence and Volatility in the information set
2	Conditions IRF estimation on recession periods only
3	Use the CFCI to measure Financial Conditions
4	Uses the S_{plus} function to measure spikes in the EBP
5	Uses the S_{diff} function to measure spikes in the EBP
6	Uses a recursive scheme to identify shocks to the EBP

Notes: Confidence is the Michigan Sentiment Index. Volatility is the measure of Uncertainty computed by Jurado, Ludvigson and Ng (2015), see Jurado, Kyle, Sydney C. Ludvigson, and Serena Ng. 2015. "Measuring Uncertainty." American Economic Review, 105(3): 1177-1216. Recession periods are defined as periods in which output at time t is lower than output a year before. The CFCI is the Chicago Financial Condition Index. The function S_{diff} is defined as $x_t^+(j) = (x_t - x_{t-h})I[x_t - x_{t-h} > 0]$. The function S_{plus} is defined as $x_t^+ = x_t I[x_t > 0]$.

Table 3: PREDICTIVE POWER OF UNEMPLOYMENT NEWS FOR ECONOMIC ACTIVITY

Forecast horizon	Without Controls			With Controls			
	6	12	18	6	12	18	
Variable	<i>t</i> -statistics			<i>t</i> -statistics			
Unemployment	1.69	1.93	2.49	Unemployment	1.33	1.64	2.10
IP	-1.39	-2.05	-2.79	IP	-0.88	-1.57	-2.16
Employment	-2.05	-2.27	-2.86	Employment	-1.63	-1.64	-1.83

Notes: The table reports the t-statistics of the coefficients related to $Unews_t$ in equation $\nabla^h Y_{t+h} = a(L)\Delta Y_t + \Gamma'x_t + \beta U_{t-1}^{news} + \epsilon_{t+h}$ where Y is the Unemployment Rate, the log of Industrial Production or the log of Employment and x_t is a vector of controls containing the usual variables in the Gilchrist and Zakrajsek (2012) framework (i.e., real fed funds, the term spread and the predicted GZ spread). Standard errors are computed using the Newey-West Estimator. The sample size runs from January 1999 to December 2012, $T = 162, 156, 150$, respectively. In bold we highlight t-statistics that denote coefficients significantly different from zero at least at the 10% confidence level.

Table 4: REACTION OF EBP TO BAD AND GOOD UNEMPLOYMENT NEWS

Variable	Without Controls		With Controls	
	Coefficient	<i>p</i> -value	Coefficient	<i>p</i> -value
constant	-0.06	0.22	-0.30	0.00
EBP_{t-1}	0.69	0.00	0.69	0.00
EBP_{t-2}	0.20	0.05	0.19	0.05
$Good\ News_{t-1}$	0.15	0.56	-0.03	0.89
$Bad\ News_{t-1}$	1.31	0.03	1.44	0.02

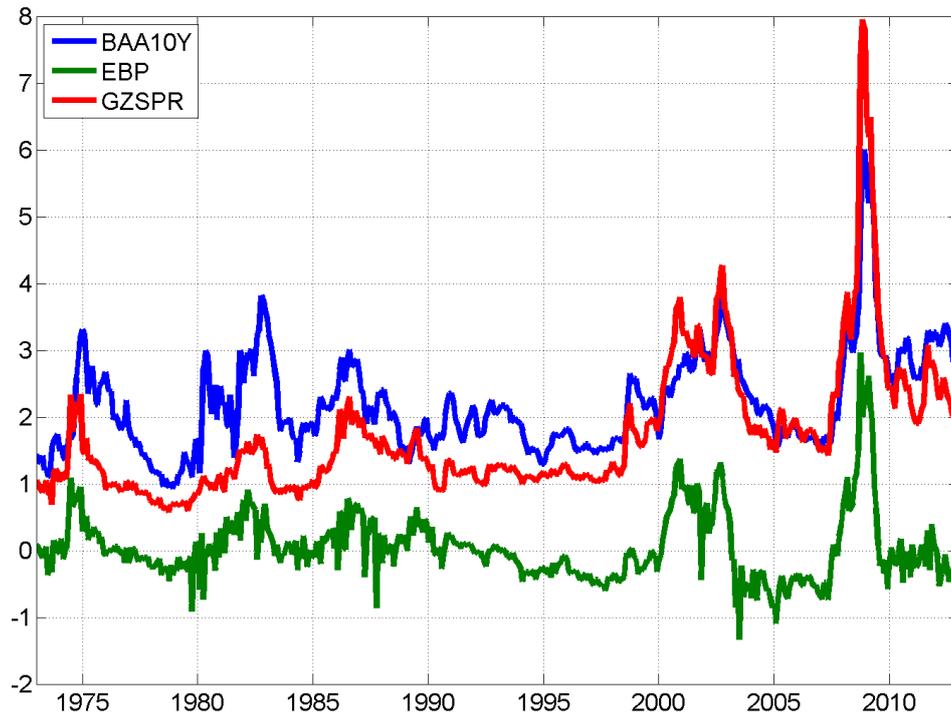
Notes: The table reports the results relative to the equation $EBP_t = \alpha + \phi(L)EBP_t + \beta(L)^{bad}U_t^{news}I(U_t^{news} > 0) + \beta(L)^{good}U_t^{news}I(U_t^{news} \leq 0) + \Gamma'x_t + u_t$. x_t is a vector of controls containing the usual variables in the Gilchrist and Zakrajsek (2012) framework (i.e., real fed funds, the term spread and the predicted GZ spread, which are not reported here). The sample size runs from January 1999 to December 2012, $T = 167$ after adjustments. Bold figures denote coefficients significantly different from zero at conventional confidence levels.

Table 5: UNEMPLOYMENT AND MONETARY POLICY NEWS

Variable	Coefficient	<i>p</i> -value.
constant	-0.01	0.04
$News_t$	-0.02	0.42
$News_{t-1}$	-0.01	0.41

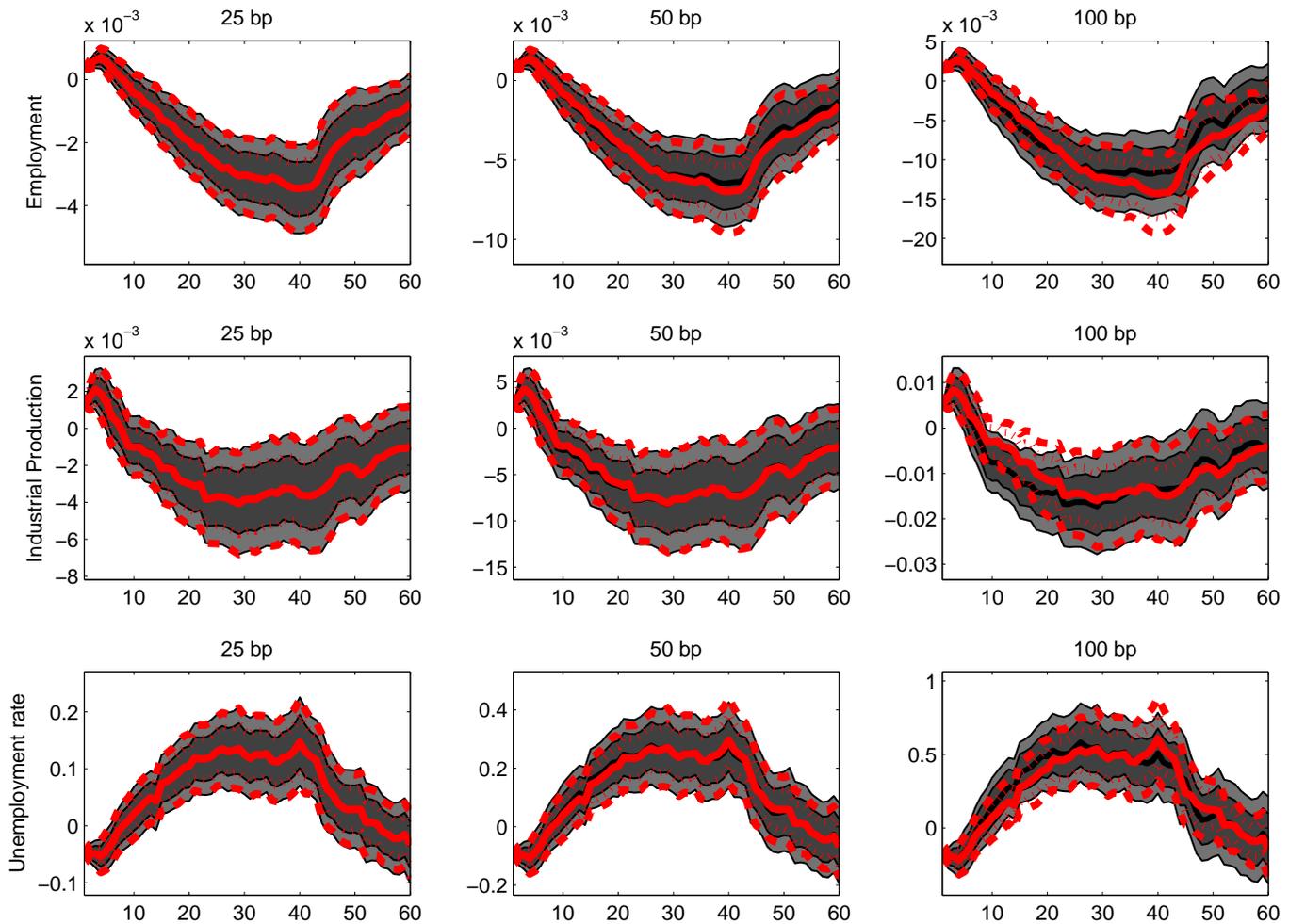
Notes: The table reports the results relative to the equation $U_t^{news} = \alpha + \beta MP_{surprise} + u_t$. The sample size runs from January 1999 to December 2012, $T = 154$ after adjustments. Bold figures denote coefficients significantly different from zero at conventional confidence levels.

Figure 1: GZ SPREAD, MOODY'S BAA-AAA SPREAD AND EXCESS BOND PREMIUM



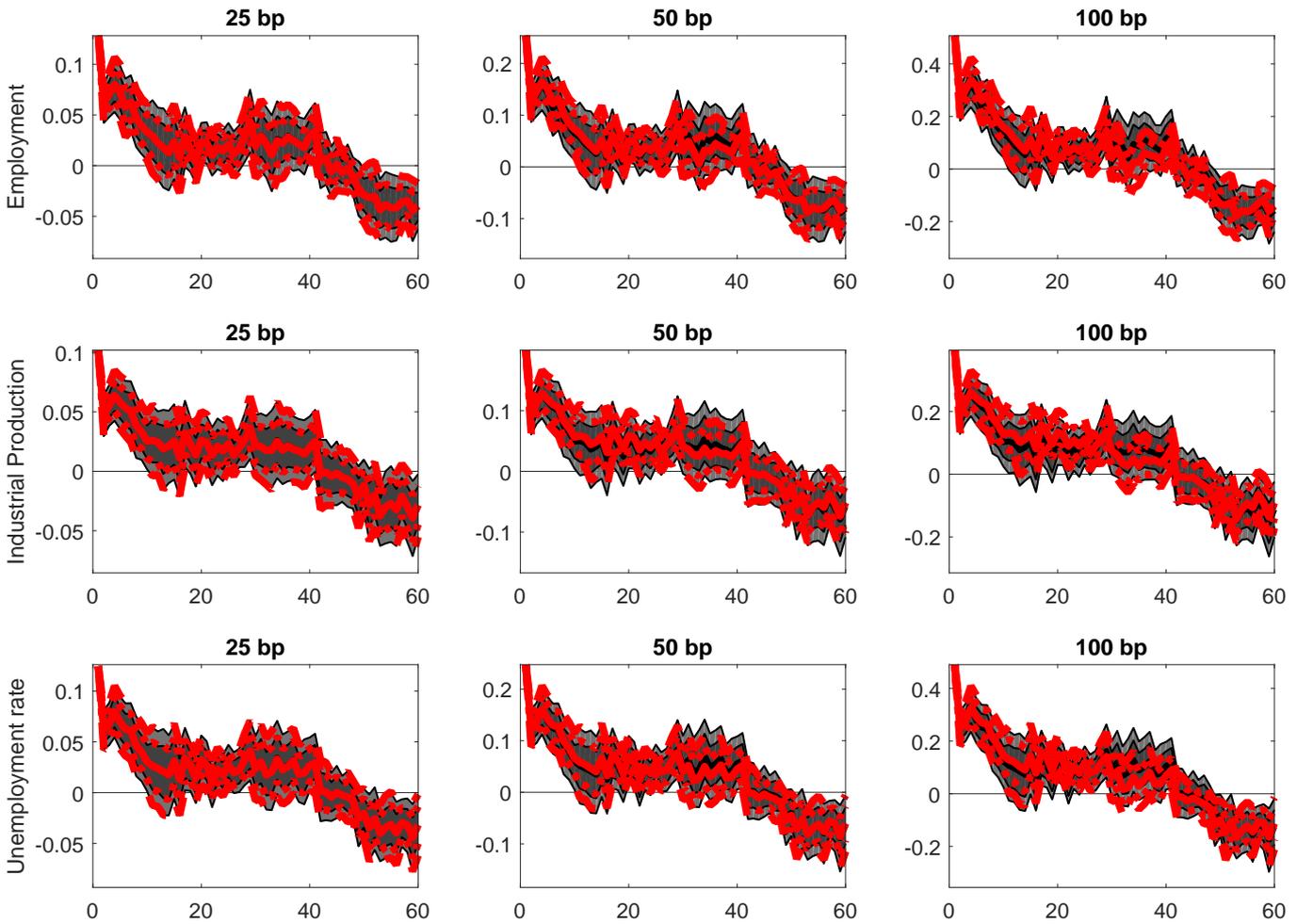
Notes: The figure compares three different indicators of tensions in credit markets, i.e., the GZ spread (red line), the Moody's spread between Seasoned BAA and AAA Corporate Bond Yield (blue line) and the EBP (green line). Sample is 1973:01 - 2012:12.

Figure 2: THE IMPACT OF MONETARY SHOCKS ON ECONOMIC ACTIVITY



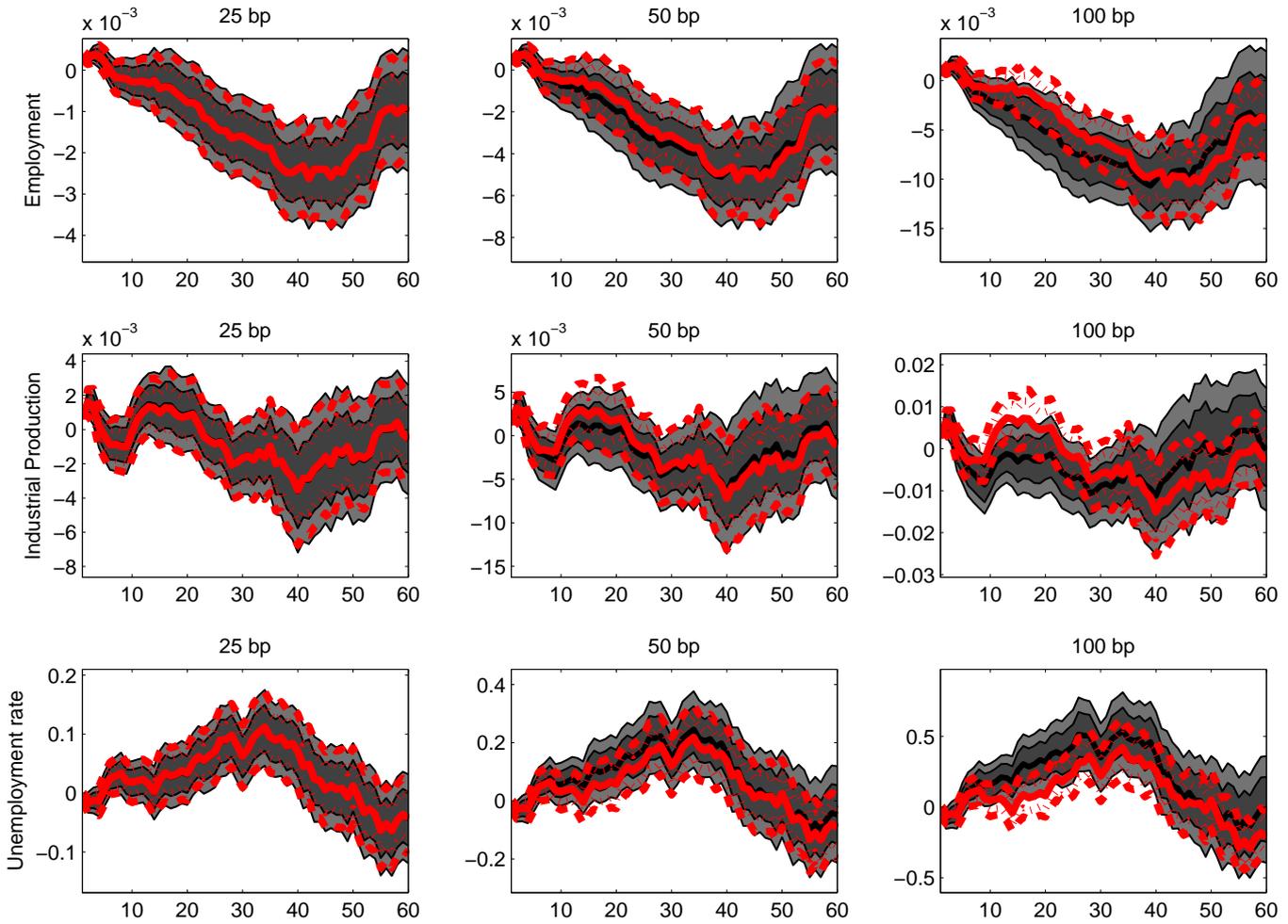
Notes: The figure shows the response of economic activity to contractionary monetary policy shocks in the USA. The rows refer to three alternative models where economic activity is measured respectively by employment, industrial production and unemployment. For each model, the columns report the responses to shocks of 25, 50 and 100 basis points. The shocks are identified using financial market data (Gertler and Karadi, 2015) and the responses are calculated by local projections (Jordà, 2005), allowing for a non-linear response of output to bond spreads. Black lines represent the estimated median responses to a monetary tightening together with their 68% and 90% confidence bands (grey shaded area). Red lines represent the estimated median responses to a monetary easing, again with their 68% and 90% confidence bands (dotted and dashed red lines). The responses to a monetary easing are multiplied by minus one. Here we are conditioning on the whole history of the shocked variable and the local peak is computed over a 12 months horizon. Sample is 1973:01- 2012:12.

Figure 3: THE IMPACT OF MONETARY SHOCKS ON EBP



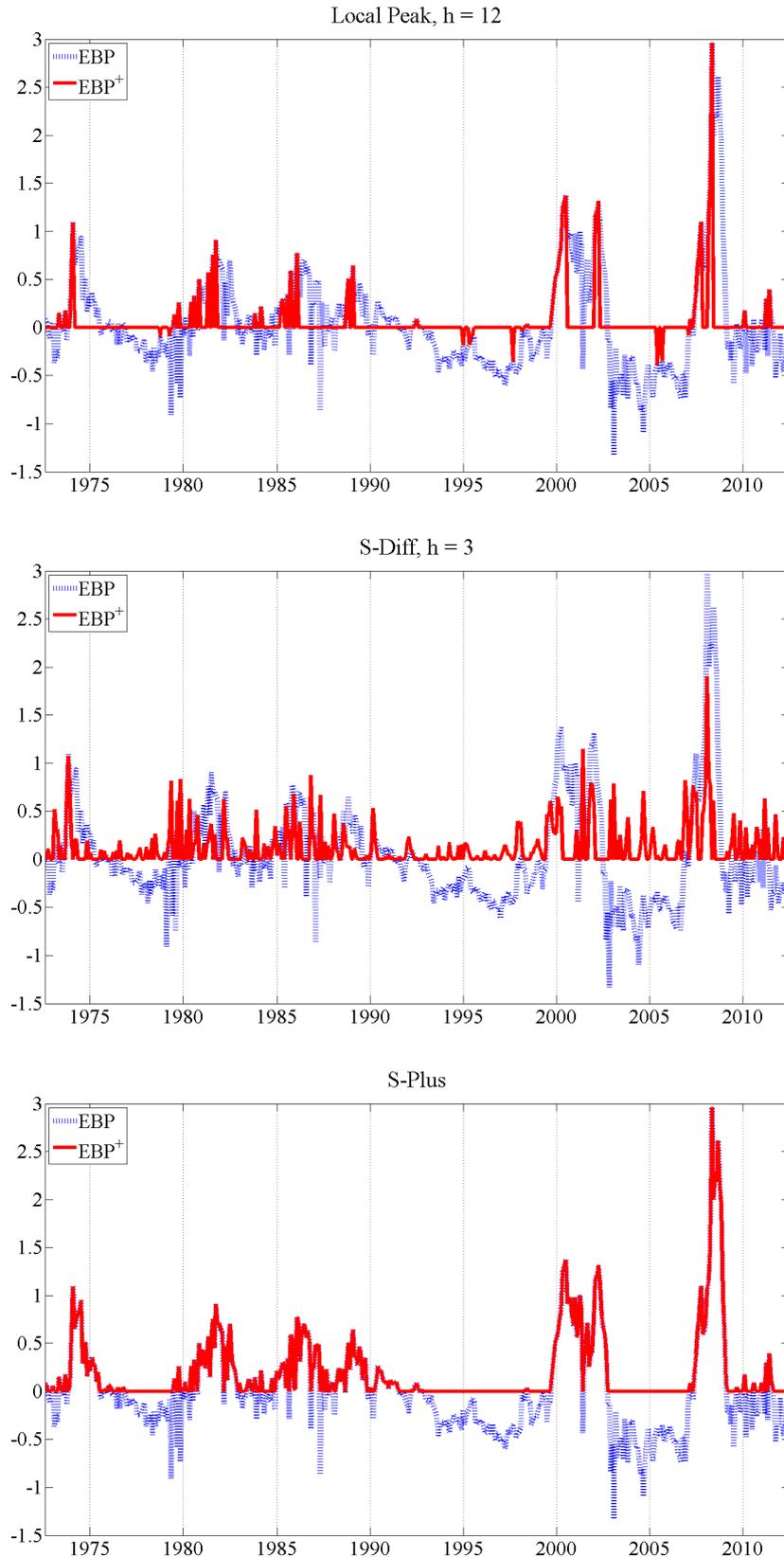
Notes: The figure shows the response of the Excess Bond Premium to monetary policy shocks in the USA. The rows refer to three alternative models where economic activity is measured respectively by employment, industrial production and unemployment. For each model, the columns report the responses to shocks of 25, 50 and 100 basis points. The shocks are identified using financial market data Gertler and Karadi (2015) and the responses are calculated by local projections (Jordà, 2005) allowing for a non-linear response of bond spreads to interest rates. Black lines represent the estimated median responses to a monetary tightening together with their 68% and 90% confidence bands (grey shaded area). Red lines represent the estimated median responses to a monetary easing, again with their 68% and 90% confidence bands (dotted and dashed red lines). The responses to a monetary easing are multiplied by minus one. Here we are conditioning on the whole history of the shocked variable and the local peak is computed over a 12 months horizon. Sample is 1973:01- 2012:12.

Figure 4: THE IMPACT OF MONETARY SHOCKS ON ECONOMIC ACTIVITY: ACCOUNTING FOR UNCERTAINTY AND CONFIDENCE



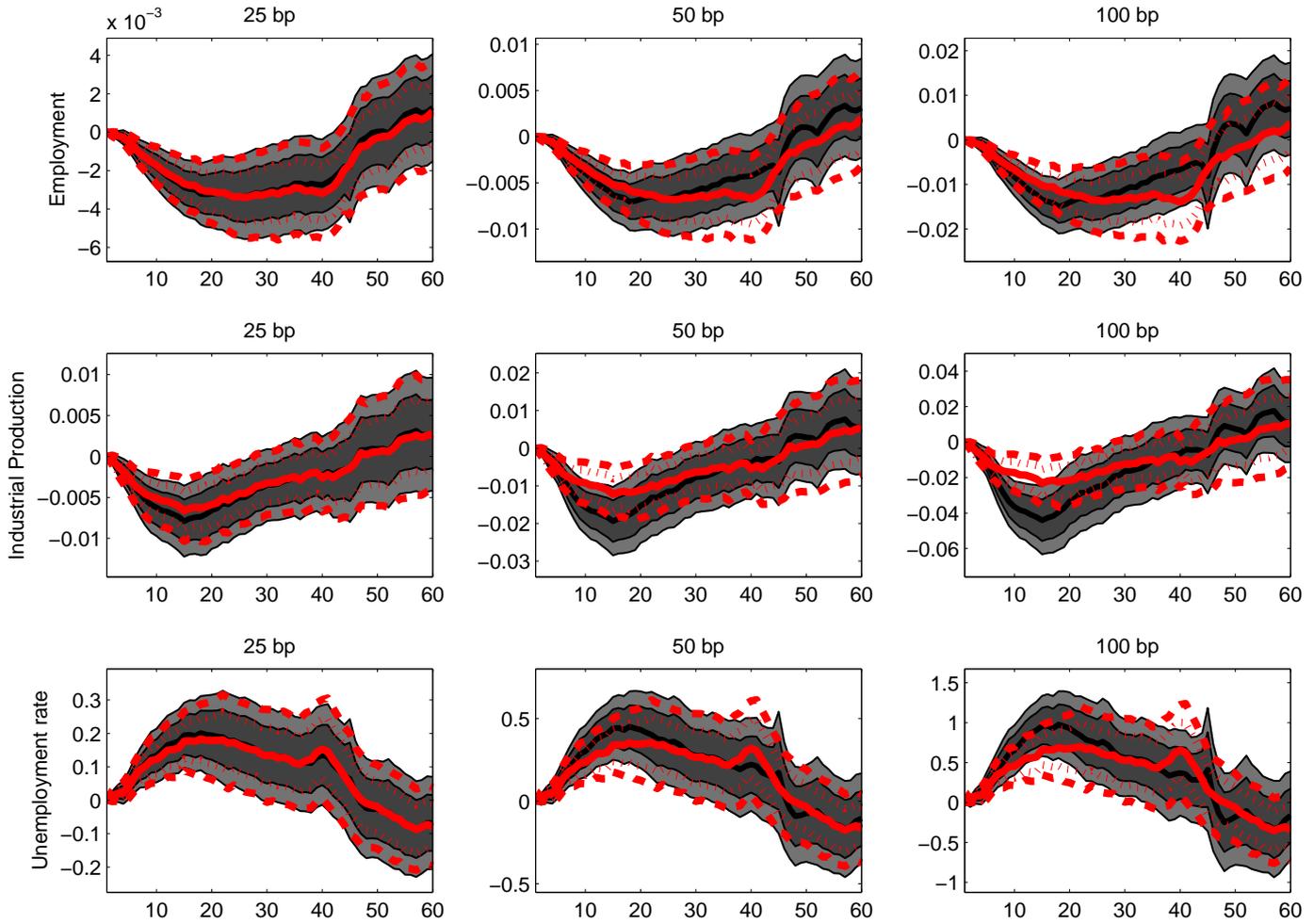
Notes: The black line represents the estimated median impulse response to a positive shock, i.e. an *increase* of the spreads, together with its 68% (the dark grey shaded area) and its 90% (the light grey shaded area). The red line represents the estimated median impulse response to a negative shock, i.e., a *decrease* in the spreads, together with its 68% (dotted red lines) and its 90% (dashed red lines). From left panel to the right increasing sizes of the shocks are plotted. Here we are conditioning on a given point in time of the shocked variable and the net increase is computed over a 12 months horizon. Sample is 1973:01 - 2012:12.

Figure 5: EBP AND ITS TRANSFORMATIONS CAPTURING ASYMMETRIES



Notes: Each subfigure plots the EBP (x_t) and the related asymmetric indicator built at horizon j ($x_t^+(j)$). The top panel shows the EBP and the local peak transformation, defined as $x_t^+(j) = x_t I[x_t > \max(x_{t-1}, x_{t-2}, \dots, x_{t-j})]$ with $j = 12$. The middle panel shows the EBP and the S_{diff} transformation, defined as $x_t^+(j) = (x_t - x_{t-h}) I[x_t - x_{t-h} > 0]$. The bottom panel shows the EBP and the S-minplus transformation defined as $x_t^+ = x_t I[x_t > 0]$

Figure 6: THE IMPACT OF AN EBP (CREDIT SPREAD) SHOCK ON ECONOMIC ACTIVITY



Notes: The figure shows the response of economic activity to monetary policy shocks in the USA. The rows refer to three alternative models where economic activity is measured respectively by employment, industrial production and unemployment. For each model, the columns report the responses to shocks of 25, 50 and 100 basis points. The shocks are identified using a standard recursive scheme (Cholesky identification) and the responses are calculated by local projections (Jordà, 2005), allowing for a non-linear response of output to bond spreads. Black lines represent the estimated median responses to a monetary tightening together with their 68% and 90% confidence bands (grey shaded area). Red lines represent the estimated median responses to a monetary easing, again with their 68% and 90% confidence bands (dotted and dashed red lines). The responses to a monetary easing are multiplied by minus one. Here we are conditioning on the whole history of the shocked variable and the local peak is computed over a 12 months horizon. Sample is 1973:01- 2012:12.

Appendix

A Monetary policy trade-off in a two-period economy

In Section 2.1 of the paper we examine a stylized two-period economy described by the following equations:

$$\begin{aligned}y_t &= \gamma \Delta s_t + \xi \Delta s_t I_{\Delta s_t > 0} + e_t \\ \Delta s_t &= -(1 - \rho) s_{t-1} + i_t\end{aligned}$$

The output gap y_t is affected by a random disturbance e_t and by the variation in credit spreads relative to the previous period, Δs_t . The impact of the spreads on economic activity is negative ($\gamma < 0$) and potentially non-linear ($\xi \leq 0$). The spread equation comes from a simple AR(1) process $s_t = (1 - \rho)s^* + \rho s_{t-1} + i_t$ with $s^* = 0$, where the equilibrium value s^* is set to zero to save notation. We consider an economy that starts off from an equilibrium situation where $y_0 = s_0 = 0$. At time 1 an exogenous shock e_1 takes place, the central bank (CB) observes it and decides whether and how to tackle it by manipulating i_1 . No actions and no further shocks take place at time 2. Conditional on the shock e_1 , the output gaps at $t = 1$ and 2 are a known function of the policy response:

$$\begin{aligned}y_1 &= \gamma i_1 + \xi i_1 I_{i_1 > 0} + e_1 \\ y_2 &= \gamma \Delta s_2 + \xi \Delta s_2 I_{\Delta s_2 > 0} = -(1 - \rho) i_1 (\gamma + \xi I_{i_1 < 0}) \quad ,\end{aligned}$$

where we used the fact that $\Delta s_2 > 0 \Leftrightarrow s_1 < 0 \Leftrightarrow u_1 < 0$, so $I_{\Delta s_2 > 0} = I_{u_1 < 0}$. In other words, given the nature of the spread equation, the non-linearity is triggered for sure in $t = 2$ if and only if the policy rate is lowered in $t = 1$. We assume that the CB discounts the future gap at a rate $\beta < 1$ and focus on a recession $e_1 < 0$ that gives the CB an incentive to implement monetary stimulus. We consider the optimal policy choice under risk neutrality and risk aversion.

Risk-neutral central bank. For the risk-neutral (*RN*) CB, the loss function is the expected (or average) output gap over the two periods, which can be written as a function of shock and policy response as follows (the time subscript can be omitted because both are dated time-1):

$$\begin{aligned}
\ell^{RN}(e, i) &= y_1 + \beta y_2 \\
&= e + \gamma i + \xi I_{i>0} - \beta(1 - \rho)(\gamma + \xi I_{i<0})i \\
&= e + \gamma i - \beta(1 - \rho)(\gamma + \xi)i
\end{aligned}$$

(The indicator function can be dropped once we focus on $e < 0$ and thus $i < 0$). The optimal policy choice can be derived by simply setting $\ell^{RN}(e, i) = 0$ and rearranging the terms:

$$i = -\frac{1}{\gamma} \left[\frac{1}{1 - \beta(1 - \rho)(1 + \xi/\gamma)} \right] e \equiv -\frac{\kappa^{RN}(\xi)}{\gamma} e,$$

where

$$\kappa^{RN}(\xi) \equiv \frac{1}{1 - \beta(1 - \rho) \left(1 + \frac{\xi}{\gamma}\right)}$$

We assume $\beta(1 - \rho)(1 + \xi/\gamma) < 1$ in order to guarantee $\kappa^{RN}(\xi) > 0$, so that $e < 0$ always implies $i < 0$.²¹ Subject to that, one can see that:

$$\begin{aligned}
i) \quad \kappa^{RN}(\xi) &\geq 1 \\
ii) \quad \kappa^{RN}(\xi) &= 1 \Leftrightarrow \rho = 1 \text{ or } \beta = 0 \\
iii) \quad \kappa_{\xi}^{RN}(\xi) &= - \left[1 - \beta(1 - \rho) \left(1 + \frac{\xi}{\gamma}\right) \right]^{-2} \frac{-\beta(1 - \rho)}{\gamma} = \frac{1}{\gamma} \frac{\beta(1 - \rho)}{\left[1 - \beta(1 - \rho) \left(1 + \frac{\xi}{\gamma}\right) \right]^2} < 0
\end{aligned}$$

These are summarised under Result (1) in the paper. With $\beta = 0$ or $\rho = 1$ the CB fully accomodates the shock, in the sense that it simply keeps the time-1 output gap constant at zero (ii). The negative time-2 gap is disregarded ($\beta = 0$) or it does not arise in the first place if the spread is random walk ($\rho = 1$). In general, the response goes beyond full accomodation (i). This multiplier effect arises because, under risk neutrality, the CB chooses a positive gap in $t = 1$ that compensates for the discounted negative gap that will materialize in $t = 2$. The emergence of a non-linearity in the transmission mechanism makes the CB even more aggressive in this respect (iii).

Risk-averse central bank Under risk aversion, the CB minimises the variance of the

²¹The condition is economically sensible – it implies that the policy rate drops (rises) after a negative (positive) shock – and not overly restrictive. It clearly holds instance if $\xi \geq \gamma$ and $\rho \leq 0.5$, as in this cases it is satisfied as long as $\beta < 1$. A smaller upper bound for β would be consistent with $t = 2$ being a shorthand for some indefinite future period.

output gap around its zero target:

$$\begin{aligned}
\ell^{RA}(e, i) &= y_1^2 + \beta y_2^2 \\
&= (e + \gamma i + \xi I_{i>0})^2 + \beta [-(1 - \rho)(\gamma + \xi I_{i<0})i]^2 \\
&= e^2 + 2e\gamma i + [\gamma^2 + \beta(1 - \rho)^2(\gamma + \xi)^2] i^2
\end{aligned}$$

The first-order condition for this problem is:²²

$$\begin{aligned}
\ell_i^{RA}(e, i) &= 2e\gamma + 2[\gamma^2 + \beta(1 - \rho)^2(\gamma + \xi)^2]i = 0 \\
i &= -\frac{\gamma}{[\gamma^2 + \beta(1 - \rho)^2(\gamma + \xi)^2]}e \\
&\equiv -\frac{\kappa^{RA}(\xi)}{\gamma}e,
\end{aligned}$$

where

$$\kappa^{RA}(\xi) \equiv \left[\frac{1}{1 + \beta(1 - \rho)^2 \left(1 + \frac{\xi}{\gamma}\right)^2} \right]$$

In this case the multiplier has the following properties:

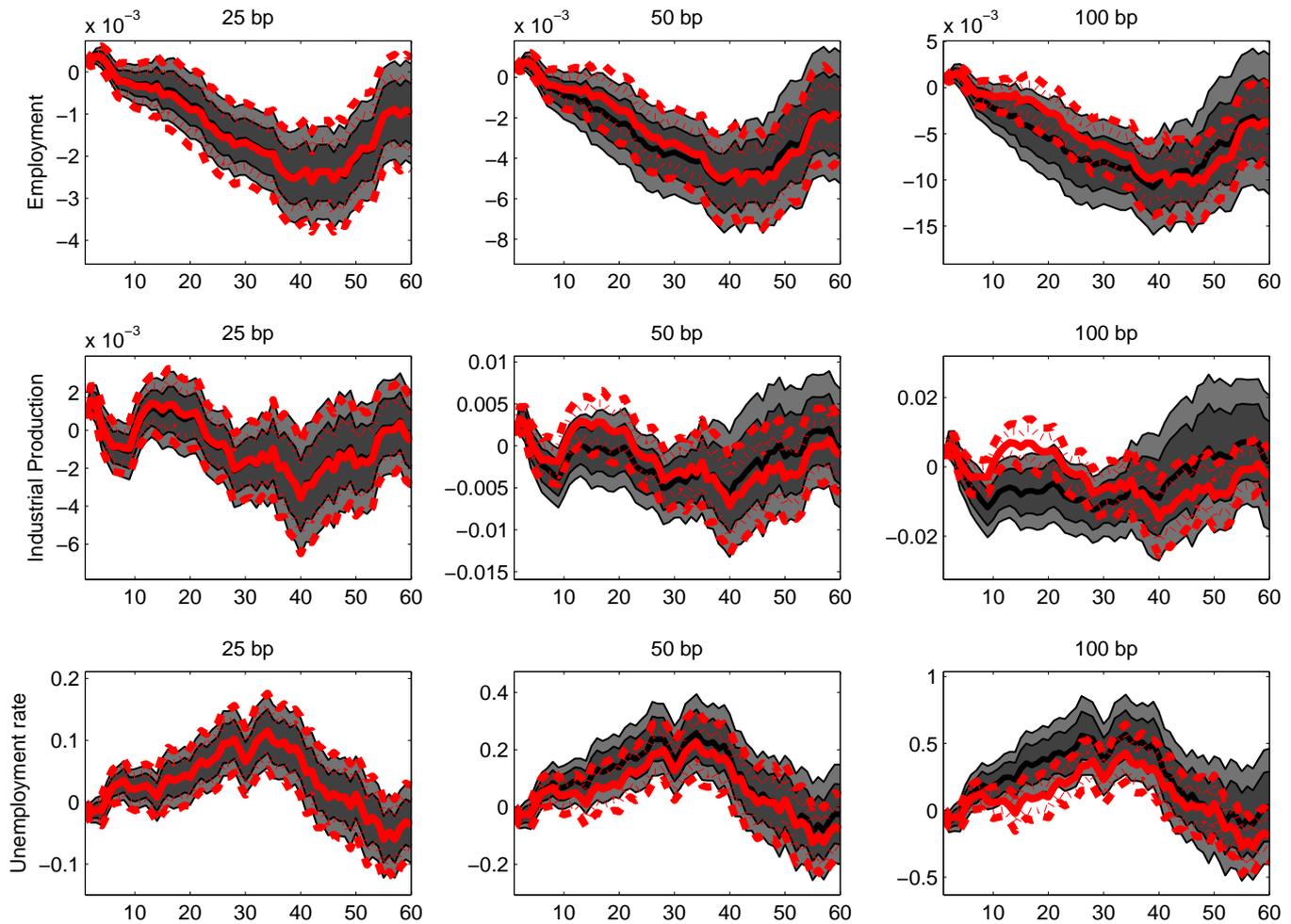
$$\begin{aligned}
i) \quad \kappa^{RA}(\xi) &\leq 1 \\
ii) \quad \kappa^{RA}(\xi) &= 1 \Leftrightarrow \rho = 1 \text{ or } \beta = 0 \\
iii) \quad \kappa_\xi^{RA}(\xi) &= -\left[1 + \beta(1 - \rho)^2 \left(1 + \frac{\xi}{\gamma}\right)^2\right]^{-2} \frac{2}{\gamma} \left(1 + \frac{\xi}{\gamma}\right) \\
&= -\frac{1}{\gamma} \frac{2\left(1 + \frac{\xi}{\gamma}\right)}{\left[1 + \beta(1 - \rho)^2 \left(1 + \frac{\xi}{\gamma}\right)^2\right]^2} > 0,
\end{aligned}$$

(where the last inequality in (iii) follows again from $\gamma < 0$ and $\xi \leq 0$). Risk aversion generally creates an attenuation effect: relative to the benchmark case of an impatient CB (or a random-walk spread), the interest rate here moves less (i, ii). That implies *a fortiori* that the risk-averse CB acts less than the risk-neutral CB examined above. Furthermore, the non-linearity works in the opposite direction compared to the risk neutral case, leading to even milder policy interventions (iii).

²²The second-order condition is satisfied so this identifies the global minimum for the loss function.

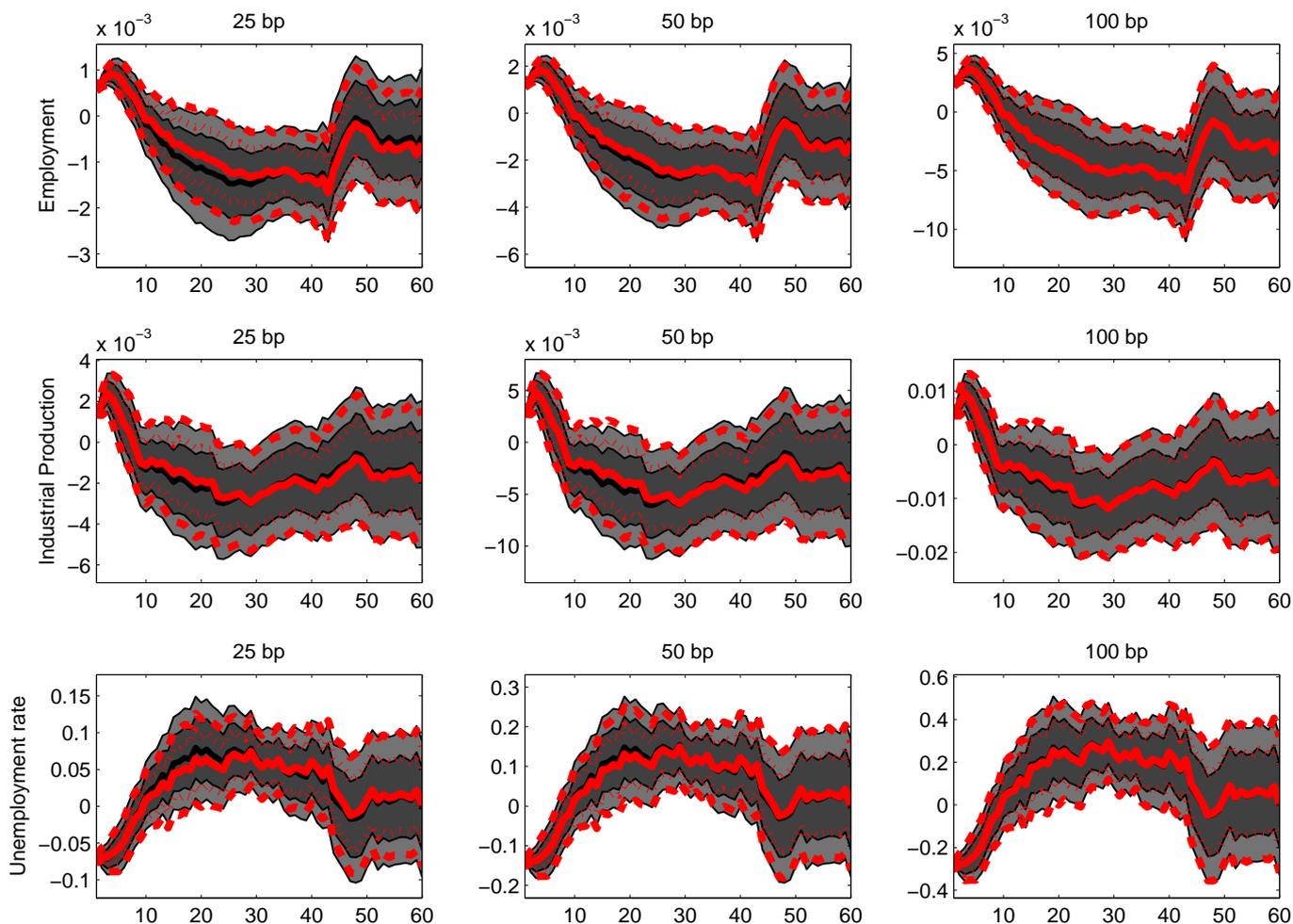
B Supplementary Figures

Figure B-1: THE IMPACT OF MONETARY SHOCKS ON ECONOMIC ACTIVITY, CONDITIONING ON RECESSIONS (SPECIFICATION 2 ON TABLE 2 OF THE PAPER)



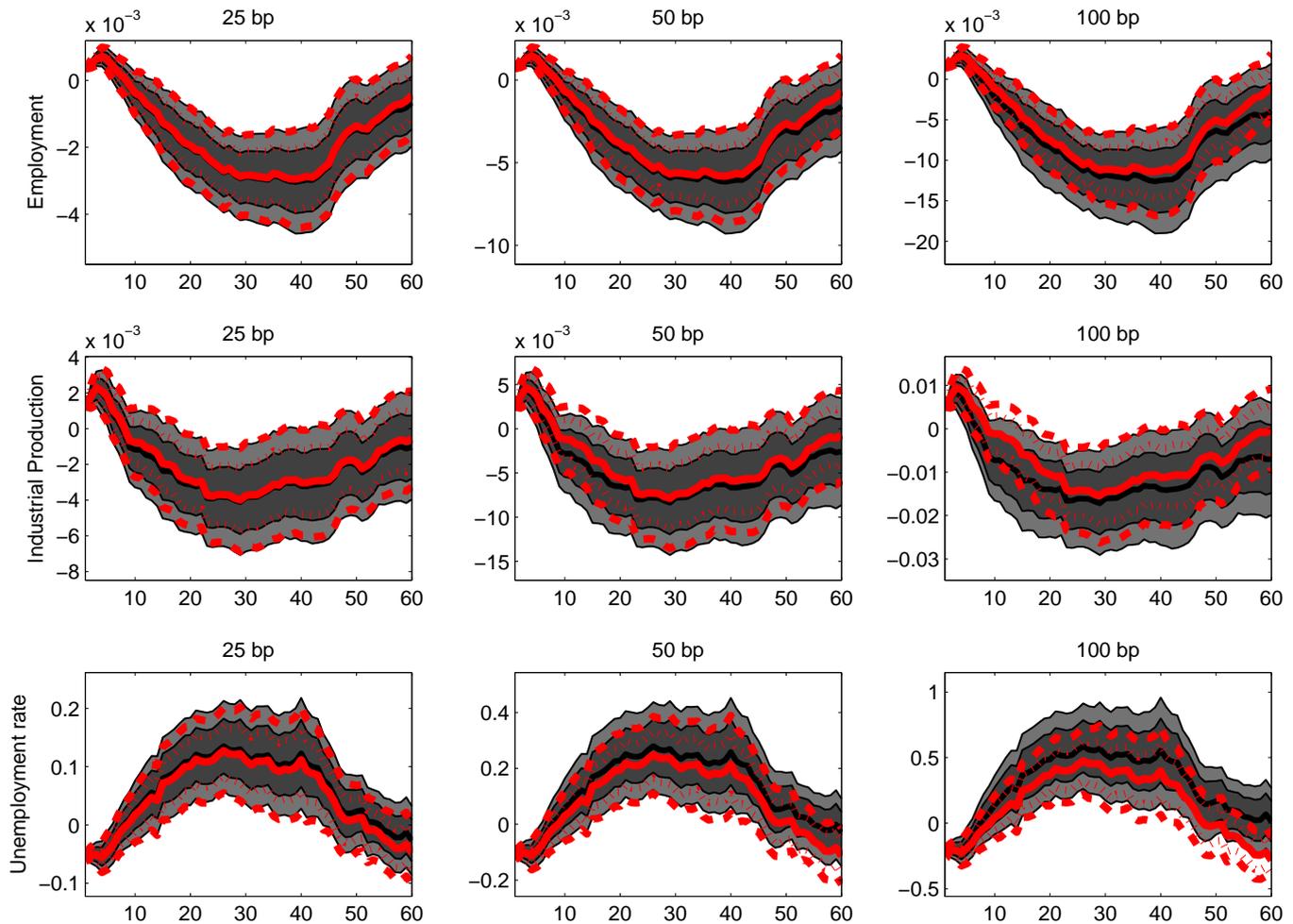
Notes: The black line represents the estimated median impulse response to a positive shock, i.e. an *increase* of the spreads, together with its 68% (the dark grey shaded area) and its 90% (the light grey shaded area). The red line represents the estimated median impulse response to a negative shock, i.e., a *decrease* in the spreads, together with its 68% (dotted red lines) and its 90% (dashed red lines). From left panel to the right increasing sizes of the shocks are plotted. Here we are conditioning on a given point in time of the shocked variable and the net increase is computed over a 12 months horizon. Sample is 1973:01 - 2012:12.

Figure B-2: THE IMPACT OF MONETARY SHOCKS ON ECONOMIC ACTIVITY: CHICAGO FCI AS FINANCIAL CONDITIONS INDICATOR (SPECIFICATION 3 IN TABLE 2 OF THE PAPER)



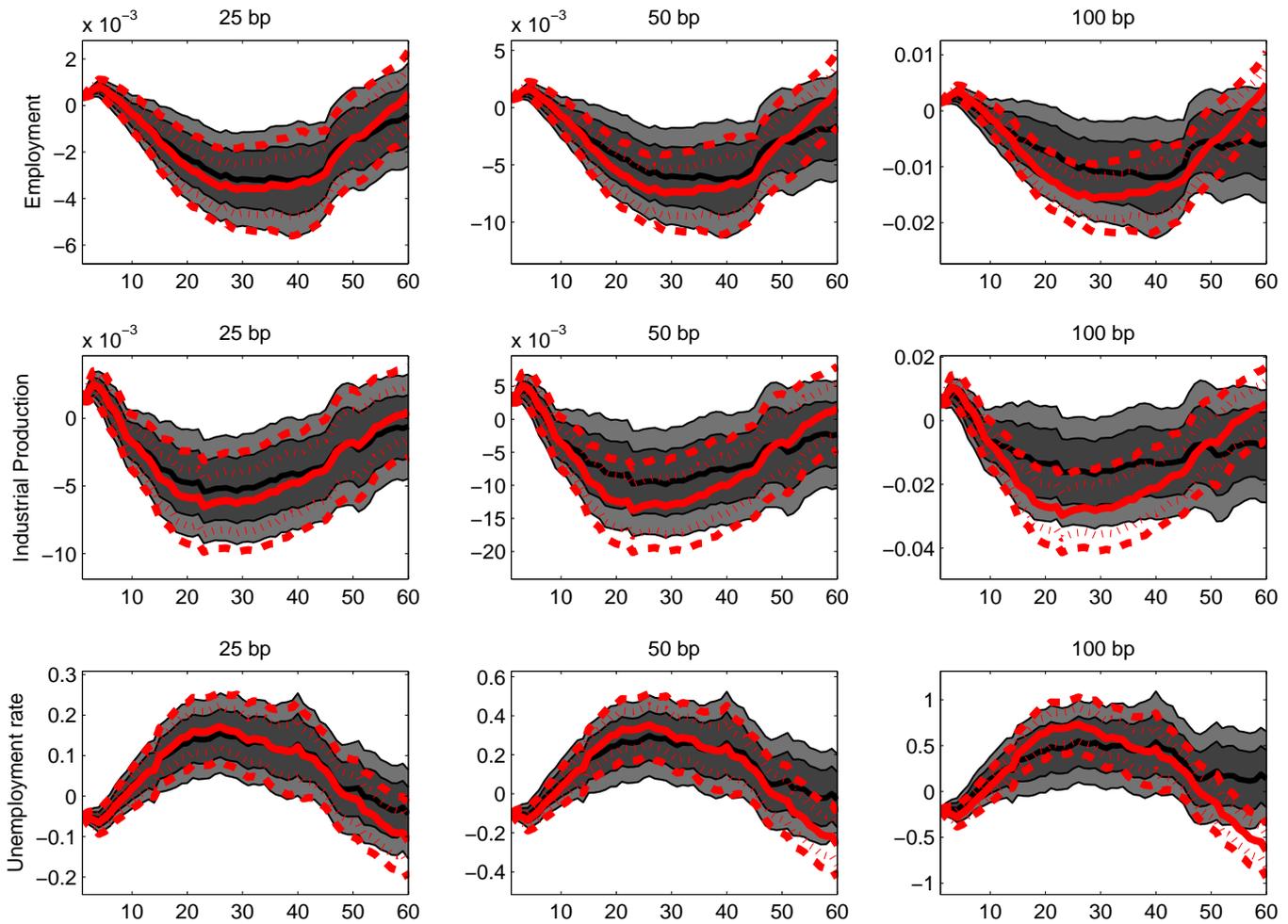
Notes: The black line represents the estimated median impulse response to a positive shock, i.e. an *increase* of the spreads, together with its 68% (the dark grey shaded area) and its 90% (the light grey shaded area). The red line represents the estimated median impulse response to a negative shock, i.e., a *decrease* in the spreads, together with its 68% (dotted red lines) and its 90% (dashed red lines). From left panel to the right increasing sizes of the shocks are plotted. Here we are conditioning on a given point in time of the shocked variable and the net increase is computed over a 12 months horizon. Sample is 1973:01 - 2012:12.

Figure B-3: THE IMPACT OF MONETARY SHOCKS ON ECONOMIC ACTIVITY, S_{diff} AS NON-LINEAR TRANSFORMATION OF EBP (SPECIFICATION 4 IN TABLE 2 OF THE PAPER)



Notes: The black line represents the estimated median impulse response to a positive shock, i.e. an *increase* of the spreads, together with its 68% (the dark grey shaded area) and its 90% (the light grey shaded area). The red line represents the estimated median impulse response to a negative shock, i.e., a *decrease* in the spreads, together with its 68% (dotted red lines) and its 90% (dashed red lines). From left panel to the right increasing sizes of the shocks are plotted. Here we are conditioning on a given point in time of the shocked variable and the net increase is computed over a 12 months horizon. Sample is 1973:01 - 2012:12.

Figure B-4: THE IMPACT OF MONETARY SHOCKS ON ECONOMIC ACTIVITY, S_{plus} AS NON-LINEAR TRANSFORMATION OF EBP (SPECIFICATION 5 IN TABLE 2 OF THE PAPER)



Notes: The black line represents the estimated median impulse response to a positive shock, i.e. an *increase* of the spreads, together with its 68% (the dark grey shaded area) and its 90% (the light grey shaded area). The red line represents the estimated median impulse response to a negative shock, i.e., a *decrease* in the spreads, together with its 68% (dotted red lines) and its 90% (dashed red lines). From left panel to the right increasing sizes of the shocks are plotted. Here we are conditioning on a given point in time of the shocked variable and the net increase is computed over a 12 months horizon. Sample is 1973:01 - 2012:12.

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