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Is Deflation Cause For Panic? Evidence from the National Banking Era*

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Abstract

This paper reexamines the traditional view that all unanticipated deflation can lead to bank panics. I identify two distinct deflationary shocks by employing a sign-restricted VAR on U.S. National Banking era with monthly data for prices, real output, and bank panics. While a negative aggregate demand shock increases the likelihood of a bank panic by $3.4\% - 8.4\%$, a positive aggregate supply shock has no significant effect. My results, therefore, align with recent theoretical work arguing that deflation's impact on banking panics also hinges on real output dynamics. Hence, not all deflation is cause for panic.

Keywords: Bank Panics, Deflation, U.S. Monetary History, Sign Restrictions
JEL Codes: E31, E32, E44, E50, N11, N21

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1 Introduction

Unanticipated deflation has long been understood to contribute to the likelihood of bank panics occurring, going back at least to the work of Fisher (1933) and his debt-deflation theory of depressions.[1] The intuition was that unanticipated deflation increases real debt burdens when debt is argeed upon in fix nominal terms, and therefore heighten the risk of bank panics.[2] However, recent theoretical work by Koenig (2013), Sheedy (2014), Azariadis et al. (2019), Bullard & DiCecio (2019), and Bullard & Singh (2020) suggests that deflation resulting from positive shocks to productivity—considered a positive aggregate supply shock—may not lead to the same outcome due to the mitigating effects of the resulting higher-than-expected real incomes. Nevertheless, despite these theoretical contributions, evidence supporting this hypothesis remains scarce, making it difficult to determine if this alternative theory has empirical support.

In this paper, I make progress on this issue by examining the relationship between deflation, output and bank panics in the United States (U.S.) using monthly data from 1868-1913. This is an ideal setting as this period featured many deflationary episodes and several well-documented bank panics, making it possible to examine these relationships empirically. It was also a period of relative institutional stability as it corresponds to the interwar (post-Civil War/pre-WWI) period of the National Banking era.[3]

I use a structural vector autoregression (VAR) employing sign restrictions, with output and prices as my endogenous variables, along with a dummy variable for months in which a

[1]See Mendoza (2006), Eggertsson & Krugman (2012), Carapella (2015), Tropeano & Vercelli (2016), Mian & Sufi (2012), Mian et al. (2013) for examples of modern research in the tradition of Fisher's debt-deflation theory. Also see King (1994), Dimand (1994), Shiller (2013) for more details on the impact Fisher (1933) has had on this line of research.

[2]The "unanticipated" nature of the deflation is important here, for if the contacting agents could have anticipated it, they presumably would have built that into their agreement. Throughout this paper, I use the term "deflation" not literally as an absolute decline in the price level (negative inflation) but as a shorthand for any lower-than-expected realized price level.

[3]Officially the beginning of the National Banking era is considered 1863 with the passage of the National Currency Act (Hendrickson 2011). However, monthly data for output is unavailable before 1868.

bank panic occurred constructed from the bank panic series of Jalil (2015). Working under the standard assumption that aggregate demand curves slope downward and that short-run aggregate supply curves slope upward, this allows me to decompose a deflationary shock into two distinct shocks – a negative aggregate demand shock in which output and prices both fall, and a positive aggregate supply shock in which real output rises while prices fall.[4] I then produce the impulse response functions from these two shocks to analyze their effects on bank panics. Beginning with Uhlig (2005), sign restrictions have become a standard method for identifying economic shocks in a VAR framework. They have also recently begun to be implemented in historical research (see, for example, Calvert Jump & Kohler (2022)).

My results provide evidence that unanticipated deflation only increases the likelihood of a bank panic occurring in the face of a collapse of aggregate demand. Depending on my specifications, these shocks increase the probability of a bank panic occurring by $3.4\% - 8.4\%$ on impact and remain significantly above zero for up-to four months after the shock. By contrast, when deflation is associated with a positive aggregate supply shock (that is with an increase in real output), there is no statistically significant impact on the likelihood of bank panics. With this, my results help build a better understanding as to when bank panics are more likely to occur and when they are not, which is important given the significant adverse effects bank panics can have on output and the real economy in general (see, for example, Grossman (1993), Jalil (2015) for evidence of this during the period of study here, Calomiris & Mason (2003) for the Great Depression, and Bernanke (2018) for the Great Recession).

This is the primary result of the paper and, to my knowledge, is the first to directly provide evidence for the hypothesis that the debt-deflation link between unexpected deflation and bank panics is not a general case scenario. There is no measurable link between unanticipated deflation and bank panics when the deflation is associated with a positive

[4] Many other works interpret shocks also of this nature as aggregate demand and supply shocks respectively as well, because of the predicted co-movements of prices and output (see Selgin (1997), Bordo & Redish (2004), Beckworth (2008), Calvert Jump & Kohler (2022)).

aggregate supply shock, that is when output and prices are moving in opposite directions. In other words, the link only exists when output and prices fall together, which implies that debt-deflation is actually less about deflation *per se* leading to bank panics, but is more a story of nominal incomes or falling aggregate demand leading to bank panics.

Given the importance of the relation of output to prices in my result, one concern may be that shocks to output could be the driving factor. However, using a similar VAR framework, Jalil (2015) has already shown that output shocks alone are not a predictor for bank panics during this period. Combined with my primary result, this further suggests that one can not simply look at price and output shocks in isolation when assessing the impact these shocks may have on bank panics, one must look at the co-movements of price and output together. Put differently, not all deflation is cause for panic.

To keep my set-up as close as possible to previous works, I obtain my primary results by following Jalil (2015) in using the Long Construction Index as my measure for output (Long 1940), the USA Annalist Wholesale Price Index as a measure of prices,[5] as well as a dummy variable for the months in which Jalil has identified a bank panic occurring. However, my general results hold when using the quarterly Real GNP and GNP deflator data from Balke & Gordon (1989) as my measure for output and prices respectively.[6] My general results are robust to a number of alternative specifications as well, including using alternative price level series, removing seasonality and trend from the data, and using alternative methodologies for generating impulse response functions.

Altogether, my findings contribute to three separate strands of literature. The first is a body of work examining the nature of debt-deflation in its relationship to central bank policy. A growing number of studies have suggested that alternative policies to price stability,

[5]I have taken the USA Annalist Wholesale Price Index from the online appendix of Jalil (2015). Jalil reports having initially found the data on globalfinancialdata.com.

[6]This exercise restricts the beginning of my sample somewhat, as the Balke & Gordon (1989) series begins in 1875. I discuss the implications of this further in Section 5

such as nominal income (NGDP) targeting, can outperform inflation targeting or price-level targeting.[7] However, stabilizing NGDP or other such proxies for aggregate demand implies a policy rule which accommodates unanticipated deflation when output growth is higher than expected, which under debt-deflation assumptions could lead to financial disintermediation and even bank panics. My findings suggest that this concern is unfounded, as deflation which occurs simultaneously with increase in real output—from a positive aggregate supply shock—has no significant effects on bank panics. Thus I bolster the findings of Selgin (1997), Sumner (2012), Koenig (2013), Sheedy (2014), Azariadis et al. (2019), Bullard & DiCecio (2019); and Bullard & Singh (2020), all who argue in favour of NGDP targeting or similar such rules in favor of price stability and who argue that supply-driven deflation is unproblematic in terms of its effects on financial intermediation and the banking sector. These works are primarily motivated by theory, as such I lend empirical support.

Koenig (2013) provides suggestive evidence for his model by looking at quarterly data on output, prices and delinquency rates from the U.S. going back to 1895. From that, a simple regression shows negative price shocks to be a predictor for increased delinquency rates only in so far as they are not offset by output growth. Both Beckworth (2007, 2019) also propose empirical tests to examine the relation between deflation and bank panics, albeit indirectly. Beckworth (2007) identifies different effects of unanticipated deflations on financial intermediation in the U.S. between $1866 - 1914$, while Beckworth (2019) studies these effects in 21 advanced economies between $2000 - 2018$. In both works, Beckworth finds evidence of deepening financial markets to be correlated with positive aggregate supply shocks and stress/strain on financial markets to be correlated with negative aggregate demand shocks. However, none of these empirical analyses look directly at the link to bank panics, as I do

[7] See, for example, Garín et al. (2016), Beckworth & Hendrickson (2020), Eggertsson et al. (2021), Chen (2021) for studies which favor NGDP targeting, Selgin (1990) favoring a productivity norm, and Sumner (1995) favouring a nominal wage target - all alternatives to inflation targeting or price-level targeting which would be accommodative to unexpected deflation under some circumstances.

in this paper.

My research also contributes to an extensive literature examining the effects of deflation in the historical record. Many studies, such as Atkeson & Kehoe (2004); Bordo & Redish (2004); Bordo et al. (2004); Beckworth (2007); and Calvert Jump & Kohler (2022) have shown that both deflation types—from the demand and the supply side—occurred historically in both North America and in Europe. In fact, it was often the case that supply driven deflation was more the norm. Yet few of these works have a strong focus on the relationship between these sources of deflation and financial intermediation in the banking sector, including the potential effects of deflation on bank panics. One exception is Jalil (2015), who has shown using similar techniques as I have employed here, that bank panics lead to a decline in both real output and the price level—the definition of a fall in aggregate demand—but was unable to find evidence in the other direction. However, his VAR was not structured to disentangle the two distinct deflationary shocks by their co-movements with output, and as such, his results miss the key findings I present here. Thus when taking our results together, we now see evidence of a feedback-loop: negative demand shocks increase the likelihood of bank panics, and bank panics cause declines in aggregate demand.

Finally, my paper contributes to the literature aiming to identify distinct economic shocks with structural VARS using historical data. Traditionally, VARs to analyze this time period have used long-run or short-run restrictions following Blanchard & Quah (1989) and Gali (1992) (for example, Keating (1996), Bordo & Redish (2004), Beckworth (2007)). Yet these restrictions cannot disentangle price shocks in a way necessary to test some of the theories discussed here. Thus my implementation of sign restrictions for for testing the effects on supply and demand shocks on bank panics is a novel innovation of this paper.

Sign restrictions are most commonly employed in the literature surrounding current policy debates (see, for example, Dedola & Neri (2007); Scholl & Uhlig (2008); Dungey & Fry (2009); and Antolín-Díaz & Rubio-Ramírez (2018)). However, my paper is one of the first to apply this method in an economic history setting as an alternative identification strategy to long-run restrictions to identify such shocks. The only other work to employ similar methods, to my knowledge, is Calvert Jump & Kohler (2022), who also use sign restrictions to identify negative aggregate demand shocks and positive aggregate supply shocks using the same strategy as employed here. However, they do not incorporate bank panics in the VAR and thus do not discuss how various shocks are likely to effect financial intermediation. Furthermore, they are only concerned with the United Kingdom (U.K.), while my study focuses on the U.S. which, for the reasons to be outlined in Section 2, is a somewhat idiosyncratic case in terms of the prevailing institutions, and as such warrants studies of its own.

The remainder of this paper will go as follows. Section 2 provides a brief historical background of the period of study, while Section 3 presents the data to be used in my analysis. Section 4 outlines the sign-restricted VAR framework used to estimate and identify the effects of unanticipated deflation on the likelihood of banking panics. Next, I present my results and give an interpretation in Section 5, along with the results of alternative specifications as robustness checks. Lastly, Section 6 summarizes and concludes while discussing the relevance of this work for today.

2 Economic Conditions during the National Banking Era

The post-Civil War era in the U.S marked a period of significant economic transformation. The U.S. economy rapidly diversified and became more integrated, largely driven by decreased transportation costs, increased population, and widespread technological advancements (Williamson 1974, Kim 1998, Calomiris & Carlson 2017). This period, stretching from the end of the Civil War to the onset of World War I, experienced substantial economic growth, with real GNP annually increasing by nearly 4% on average (Balke & Gordon 1989). Beckworth (2007) has pointed out that this growth was not just extensive but also intensive, marked by sustained increases in both per capita real GNP and real wages. As a result, living standards for the average American improved noticeably.[8] In fact, the U.S. economy during this period outperformed the post-1913 era according to macroeconomic indicators including real GDP growth and price level volatility (Hogan 2015).

However, juxtaposed against this backdrop was a banking system challenged by the limitations imposed before the end of the War, by the National Bank Act of 1864.[9] While the Act aimed to create a new national banking system, its implementation inadvertently constrained the banking sector. Under this system, nationally chartered banks could issue national (as opposed to state) notes backed by either gold reserves or securities issued by the U.S. treasury department (Grossman 1993). While the previously chartered state banks were allowed to continue operations, a 10% tax was placed on state-issued bank notes during the Civil War. This severely limited the ability of the state-chartered banks to issue currency and caused many state chartered banks to permanently close their doors (Jaremski 2013, Selgin 2000).[10] This, combined with binding reserve requirements, caused a relatively inelastic and slow growing money supply (Champ et al. 1996, Selgin et al. 2012).

Additionally, the U.S. was simultaneously dealing with the monetary arrangements set-up in wartime. During the Civil War, the U.S. government had suspended gold convertibility and begun circulating "Greenbacks", a fiat legal tender.[11]. Post-war, there was a strong push to return to the gold standard at the pre-war parity, which necessitated eliminating the "gold premium" – the discrepancy between the market and official price of gold. The Contraction Act of 1866 directed the U.S. Treasury to retire the greenbacks, and with that the monetary base was reduced by about 20% between 1865 and 1867 (Bordo et al. 2007). This

[8]In order to asses per capita GNP growth and wage growth Beckworth uses the series of Balke & Gordon (1989) and Johnston & Williamson (2022) for per capita real GNP and Balke & Gordon (1989), Johnston & Williamson (2022); and *NBER Macrohistory Database* (2022) for real wages.

[9]The First National Bank Act was called the National Currency Act and was put in place in 1863. However, it was revised a year later as the National Bank Act of 1864 (Hendrickson 2011, Jaremski 2013).

[10]Technically, the tax was part of a complementary Act – the Revenue Act of 1865.

[11]For more on the U.S. financial system during the civil war, see Hammond (1970).

policy persisted, albeit more gradually, through most of the 1870s and led to the successful resumption of gold convertibility on January 1, 1879.

The system created by the National Bank Act combined with the sustained reductions in the monetary from the Contraction Act prior to the 1880's, contributed to a falling price level, which before 1896 appeared to be the norm, with average annual price level declines of 2.1%.[12] This may not seem like a large number on its own, however, it amounts to approximately a 44% total decline in the price level from 1869-1896 (Balke & Gordon 1989).[13] Because deflation was secular throughout much of the National Banking era it is possible at least some of it was anticipated.[14] It therefore may not have unexpectedly risen the real debt-burdens of borrowers, which is the first step in how debt deflation is thought to increase the probability of bank panics.[15]

Additionally, among other restrictions, the National Bank Act also prevented banks from opening more than a single branch (a practice known as "unit banking"), ensuring a fractured banking system (Calomiris & Carlson 2017). The fractured nature of unit banking likely increased the fragility of the financial system, making banks less diversified and, therefore, more prone to risk (Wicker 2000, Carlson 2005). Bank panics never occurred in countries like Canada, where branching was legal but saw similar declining price level trends during this period (Bordo et al. 2015, Cutsinger & Pender 2023). Yet during this period in the U.S., Jalil (2015) has identified episodes of eleven bank panics (three major and eight minor).

[12]Bordo & Redish (2004) attribute the reduction in secular deflation after 1896 to a series of significant gold discoveries in Australia, California and the Yukon. Despite the banks' inability to increase the broad money supply with fixed base money due to reserve requirements, the new inflow of base money (*i.e.* gold/reserves) also allowed for greater expansion of the broad money supply.

[13]While the Civil War ended in 1865, annual GNP and the Implicit Price Deflator data from Balke & Gordon (1989) does not begin until 1869.

[14]See Calomiris (1988) for more on deflation expectations prior to the reestablishment of the gold standard in 1879.

[15]I account for this in Section 5 by using a Hamilton filter to remove a trend from my output data in an attempt to remove any anticipated price level changes. Trying to estimate what changes to macro variables were anticipated by households and what changes were not is notoriously difficult. I therefore devote a portion of that section to a longer discussion of this issue.

While debt-deflation can lead to defaults of loans and therefore declines on the asset side of banks' balance sheets. It is likely diversified banks and ones with large capital holdings (like Canada had at the time) may have been able to handle a decline in assets and without triggering bank runs. Combined with this fragility, however, unanticipated deflation could have been able to to be catalyst of bank panics.

Thus we see a complex picture where overall real growth was strong, yet year-over-year price level declined approximately 30 of the 45 years under study. Additionally, despite the positive long-run trends in growth, there were instances of depression and recessions, which were often, though not always, accompanied by bank panics. At first glance, it appears challenging to disentangle deflation's effect on either output or bank panics. The following section provides the monthly data I use in my VAR to begin to disentangle such relationships.

3 Data

To begin to understand the relationship between unanticipated deflation, output, and bank panics during this period, first a definite dating of when bank panics took place is required. For this, the creation of the Jalil (2015) bank panic series brings tremendous value, deriving and presenting a novel bank panic series for U.S. from $1825 - 1929$. Unlike previous series which were somewhat arbitrary in their identification strategies for determining when a bank panic was in fact occurring, Jalil outlines a clear rule to implement, in order to define a bank panic.[16] The rule states that a banking panic can only be said to have occurred when "there is an increase in the demand for currency relative to deposits that sparks bank runs and bank suspensions" (Jalil 2015, p.300). With this rule, Jalil was able to comb through financial and economic newspapers from throughout the U.S. to determine not only when a

[16] Before presenting his new series, Jalil gives an overview of nine previous series of banking panics for this time period, none of which agree when banking panics happened, nor do they agree even on what constitutes a banking panic.

bank panic occurred, but also where in the country in originated. As such, Jalil finds that between $1825 - 1929$ the U.S experienced 27 bank panics in total, 7 of which he concludes were "major".[17]

I create a dummy variable, b, with a 1 for any month between January 1868 and December 1913 in which a bank panic was occurring (16 months in total over a sample of 552 months), and a zero else. I am unable to use any of Jalil's panic series prior to 1868 due to lack of data for the other endogenous variables. The identified nonmajor panics tended to be local, though given the fractured nature of the banking system during this period as discussed in the previous section, it is plausible that debt-deflation could trigger panics in some areas without triggering them in others. Therefore my dummy variable includes three major bank panics and eight nonmajor, for eleven panics total.[18] The timing of the included panics is presented in Table 1.

As Jalil (2015) has been able to date the month in which bank panics begin and end, and noting that the average panic in his series lasts less than three months, using annual data appears to be too low a frequency. This precludes me from using the commonly used annual Balke & Gordon (1989) data for output and prices as their data is not available monthly. Instead I follow Jalil (2015) in using the Long Construction Index and the USA Annalist Wholesale Price Index for my output (y) and price level (p) data respectively. The Long index shows significant seasonality, and so I first remove the seasonal component using the standard X-13ARIMA-SEATS method. Both output and prices are presented in log scale in figure Figure 1.

While it may be imperfect to take a construction and wholesale price index as accurate

[17]To be classified as major, the panic must hit two requirements "(i) it spans more than one geographic unit defined as a state and its bordering states, and (ii) it appears on the front page of the newspaper. All other banking panics are classified as nonmajor" (Jalil 2015).

[18]Including both the nonmajor panics as well as all months in which panics were taking place (as opposed to merely the months in which major panics began) varies slightly from the panic dummy found in Jalil (2015), however as shown in Appendix A.2, excluding all months except those in which a major panic began does not drastically alter my results.

Table 1: Major and nonmajor Bank Panics in the U.S. between $1868 - 1913$

Year	Month(s)
1873	September (major)
1884	May
1890	November
1893	May – August (major)
1896	December
1899	December
1901	June – July
1903	October
1905	December
1907	October – November (major)
1908	January

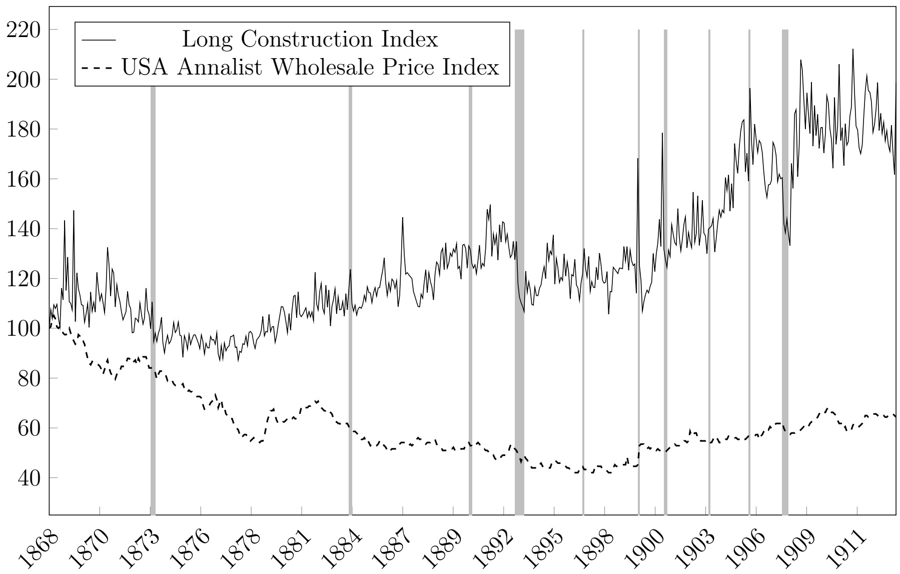
Source: Jalil (2015)

measures of overall real economic output and the general price level respectively, this data remains the best proxies available at the monthly frequency and also keeps the inputs of my VAR identical to Jalil's, implying the any differences in results must be from my decomposition of the two sources of unanticipated deflations. Jalil too recognizes this problem, but notes that construction was one of the leading investment goods industries during the National Banking era, making the Long Construction Index an reasonable proxy for real economic activity, and he also notes that Grossman (1993) has made similar arguments in favour of using this data series.[19]

With the data to be used outlined, the following section presents my structural VAR and I discuss as to how sign restrictions are imposed to create my estimations.

[19] See Appendix A.3 for VAR results using alternative available monthly price data. See Appendix A.1 where I further argue for the appropriateness of using of the Long Index as a proxy for output by checking its correlation against the Adjusted Miron-Romer Index of Industrial Production (Miron & Romer 1990), a well-know broad index often used as a proxy for output in other research.

Figure 1: Output and Prices Jan 1868 - Dec 1913 (Jan 1868 = 100)



Source: Long (1940) for construction, and Jalil (2015) for prices and bank panics.

Note: Shaded areas represent months in which bank panics occurred.

4 Methodology

Discussions of debt-deflation, as outlined in Section 1, have recently been framed in terms of the interaction between prices and output. It has been theorized that when output moves with prices during a negative aggregate demand shock, this can lead to increased debt burdens of borrowers and even bank panics, which is inline with traditional debt-deflation views. However, when output rises, offsetting falling prices during a positive aggregate supply shock, the resulting higher-than-expected real incomes will offset the rising real debt burdens. Therefore, this is not predicted to increase the likelihood of bank panics (Selgin 1997, Sumner 2012, Koenig 2013, Sheedy 2014, Azariadis et al. 2019, Bullard & DiCecio 2019, Bullard & Singh 2020).

To analyze these predictions during the National Banking era, a sign-restricted VAR is estimated where the shocks are decomposed into aggregate demand and aggregate supply shocks by the co-movements of prices and output based on standard macro assumptions of upward sloping supply curves and downward sloping demand curves. To see how this is implemented, first consider the following reduced for VAR:

$$X_t = \sum_{\tau=1}^{12} B_\tau X_{t-\tau} + u_t \ , \tag{1}$$

where where the constant term is suppressed for notational convenience.

Here X_t is a vector of the three endogenous variables under study as discussed in above in Section 3, therefore $X_t = [y_t \ p_t \ b_t]'$. The τ variable counts a discrete time interval which is chosen based on the data being used, in this case monthly, making the lags equal to 12 months or one year, following Jalil (2015). B_1, B_2 $\ldots$, B_{12} are then the 3×3 matrices of parameters. u_t is the 3 dimensional vector of errors assumed to be temporally uncorrelated, which can be represented by $\mathbb{E}\left[u_t u_{t-\tau}'\right] = 0$ for $\tau \neq 0$. Furthermore I assume $\mathbb{E}\left[u_t u_t'\right] = \Omega$, where Ω is the 3×3 variance-covariance matrix, implying that u_t follows a zero mean white noise process such that $u_t \sim (0, \Omega)$.

In order to structure (1) and impose the desired sign restrictions, first two new variables can be defined, "C" and "ε_t" such that $u_t = C\varepsilon_t$ where ε_t is a $k = 3$ dimensional vector of normalized "structural shocks" such that $\varepsilon_t \sim^{iid} (0, I)$ and $\mathbb{E}[\varepsilon\varepsilon'] = I$. C is a 3×3 matrix and assumed to be invertible. Note that from the assumptions on Ω, u_t, and ε_t, it's implied that:

$$\Omega = \mathbb{E}[u_t u_t'] = C\mathbb{E}[\varepsilon_t \varepsilon_t']C' = CC' \ . \tag{2}$$

Thus, (1) can be re-written as:

$$X_t = \sum_{\tau=1}^{12} B_\tau X_{t-\tau} + C\varepsilon_t \ . \tag{3}$$

To see clearly how (3) explicitly shows the structure of the model, it can be rearranged as follows:

$$C^{-1}X_t = \sum_{\tau=1}^{12} C^{-1}B_\tau X_{t-\tau} + \varepsilon_t \ , \tag{4}$$

$$\beta_0 X_t = \sum_{\tau=1}^{12} \beta_\tau X_{t-\tau} + \varepsilon_t \ , \tag{5}$$

or

$$\beta(L)X_t = \varepsilon_t, \tag{6}$$

where $\beta_0 = C^{-1}$, $\beta_\tau = C^{-1}B_\tau$ and $\beta(L)X_t = \beta_0 - \beta_1 L - \beta_2 L^2 \ldots \beta_{12} L^{12}$. Presented in the forms of (5) and (6), it can now clearly be seen how the elements of β_0 define the contemporaneous relationships between the elements of X_t, and ε_t as a term on its own represents the structural shocks of the model.

In order to estimate this model, B_1 through B_{12} as well as Ω can be obtained using standard OLS techniques. However, given that C has 3 degrees of freedom, obtaining an estimate for C (and therefore β_0 through β_{12}) from Ω requires additional assumptions. I impose that the elements of C have the following signs based on the discussion above:

$$\underbrace{\begin{bmatrix} u_y \\ u_p \\ u_b \end{bmatrix}}_{u_t} = \underbrace{\begin{bmatrix} - & + & * \\ - & - & * \\ * & * & * \end{bmatrix}}_{C} \underbrace{\begin{bmatrix} \varepsilon_{AD} \\ \varepsilon_{AS} \\ \varepsilon_{misc} \end{bmatrix}}_{\varepsilon_t} \tag{7}$$

Here in (7), with the imposed signs on the elements of C, ε_{AD} can now be defined as a negative aggregate demand shock, where output move in the same direction as prices, whereas ε_{AS} is defined as a positive aggregate supply shock where output moves in the opposite direction as prices. This can be conceptualised in a simple aggregate supply /

aggregate demand framework, where so long as the aggregated demand curve is assumed to slope downwards and in the short run the aggregate supply curve slopes upwards, then a shift in either curve would produce the co-movements imposed on the elements of C as shown in (7).[20] I take these assumptions as relatively unproblematic given they are the defining features of standard macro analysis, and furthermore, Calvert Jump & Kohler (2022) show explicitly how such assumptions can be derived from a workhorse three equation New Keynesian model. Calvert Jump & Kohler (2022) find these applicable restrictions to identify the same two shocks for the U.K. during a time period which overlaps with my period of study.

Elements of C marked with "$*$" are left unrestricted, which implies two things of importance. First, in neither of the two defined shocks has anything been imposed on bank panics (b). This is essential, for if a sign were to be imposed on b for either of these shocks, it would be imposing a result on the model instead of allowing it to be observed. Second, because I have only defined two shocks, the third is left unrestricted, denoted by ε_{misc}, and is therefore not interpretable economically and is of no consequence to the results under observation. Thus, having only defined two shocks in a trivariate VAR, we have left the model only partially restricted. However, following Uhlig (2005) partially sign restricted VARs have become standard practice and one only needs to analyse the shock which have been given structure based on economic theory.[21]

With this sign-identified model, the parameters of C are not point identified as would be the case in a fully defined short- or long-run restricted VAR, but are instead set identified, meaning that no unique C matrix can be estimated from Ω. Instead, a family of sign-conforming C matrices can be obtained from which to draw my analysis.

[20] By contrast, while long-run restrictions focus on the assumption of a vertical long run aggregate supply curve, here our concern is the short run as our C as this is an impact matrix.

[21] For more on the validity of partially restricted VARs in general, see Keating (1996) and Christiano et al. (1999).

In my primary results I use the Uhlig penalty-function method to obtain my family of C matrices.[22] The Uhlig penalty-function is an algorithm which penalizes draws of C when their resulting impulse response functions do not match the implied directions from the given sign restrictions and compiles a family of Cs with the smallest penalties assigned. Employing this method allows me to obtain a family of 1000 sign-conforming structural impulse responses to my identified shocks. The results and discussion of which are presented in the following section.

5 Results and Discussion

To estimate the VAR, I take the logarithm of the data for output and prices.[23] As discussed in Section 3, the dummy variable for bank panics is entered as a "1" for any month in which a major and nonmajor panic was occurring. As in Uhlig (2005), I impose that the sign restrictions hold contemporaneously as well as for 6 months, or two quarters after the shock begins.[24]

Figures 2 and 3 report the impulse response functions of the endogenous variables to a one standard deviation negative aggregate demand and a positive aggregate supply shock for 24 months, respectively. Because I have not obtained a unique estimate for C but instead 1000 sign-conforming estimates of C as discussed in Section 4, the black lines represent the median response from these 1000 draws, while the grey area represents draws between the

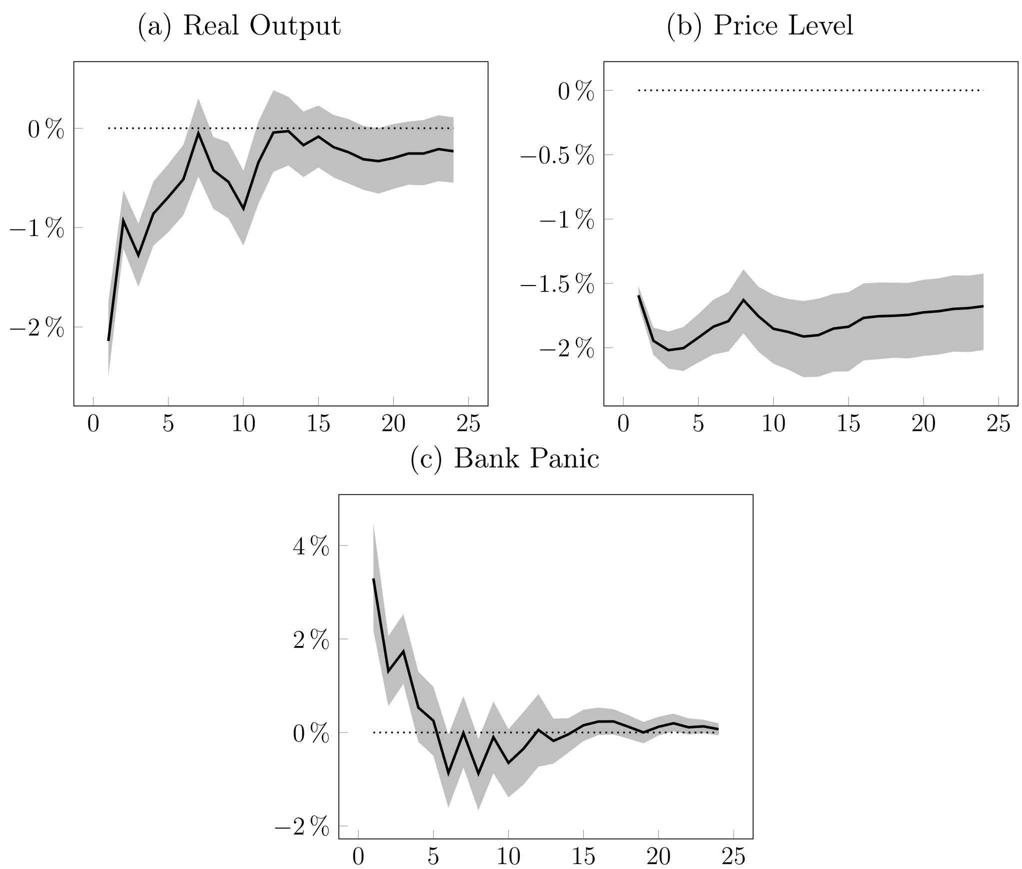
[22] Though Rubio-Ramirez et al. (2010) have proposed a rejection method as an alternative to Uhlig's penalty function. Appendix A.4 presents the impulse responses obtained using this alternative method, along with a discussion of them.

[23] While both output and prices are non-stationary in this case even in log form Sims et al. (1990), Ramey (2016); and Hamilton (2020) have established that a log-level specification is a reliable means of obtaining consistent estimates, even in the presence of variables exhibiting stochastic trends and potential cointegration. Furthermore, Elliott (1998) and Gospodinov et al. (2013) have shown that pretesting variables and imposing unit root and cointegration relationships may result in significant size distortions, both in theoretical and practical contexts. As such, the most secure approach is to estimate the VAR in log-levels, despite incorporating deterministic trends, since in this case imposing stationarity is not specifically required for identification purposes.

[24] Altering the duration for which the sign restrictions are imposed does not significantly affect my findings.

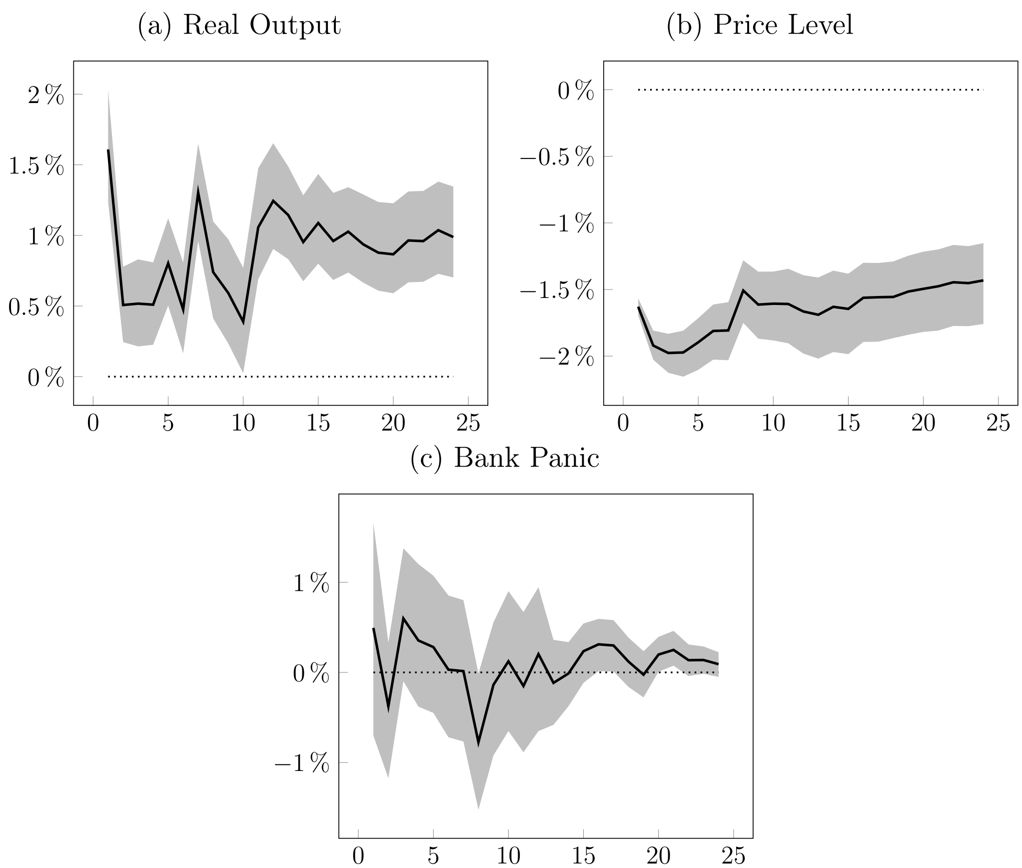
16th and 84th percentiles, which is standard practice within the sign-restriction literature (see, for example, Uhlig (2005)). Under the assumption that the 1000 draws of C are normally distributed, the grey area represents one standard deviation in either direction, which are commonly interpreted as error bands when using sign restrictions (for more on this, also see Uhlig (2005)).

Figure 2: Response to Negative AD Shock



Looking first to Figure 2, the median response of bank panics to an aggregate demand shock is for bank panics to be 3.4% more likely on impact, with the effect tapering to zero after five months. Furthermore, even the 16th percentile response is above zero for the first four months, suggesting that this result is indeed significant. Whereas in Figure 3 we see that the median response of bank panics to an aggregate supply shock is very close to zero

Figure 3: Response to Positive AS Shock



(0.5% on impact) and is not significant based on the 16th and 84th percentile responses.

This core finding supports theories outlined by Selgin (1997), Koenig (2013), Sheedy (2014), Azariadis et al. (2019), Bullard & DiCecio (2019), and Bullard & Singh (2020) suggesting that the relationship between unexpected deflation and banking crises is not a universal occurrence. While I find evidence for debt-deflation linkig deflation and bank panic on impact of a negative aggregate demand shock (when output and prices move in the same direction), any link between deflation and banking panics is insignificant when such deflation coincides with a positive aggregate supply shock, namely, when output and prices move in opposite directions.

To put these findings into a broader context, both Grossman (1993) and Jalil (2015) have provided ample evidence that banking panics during this time led to immediate large declines real economic activity. Therefore, given harmful effects bank panics cause when they do occur, any increase in likelihood of a bank panic should be noteworthy. My findings build upon Jalil (2015) by identifying a link from prices to panic in some cases where none was previously found. When we combine these new insights with Jalil's discovery that prices and output both decline after a bank panic, evidence of a negative feedback loop emerges. In this loop, contractions in demand increase the probability of banking panics, which in turn, have the potential to further collapse aggregate demand.

It should be worth noting however, that given the complex nature of the financial system and the regional institutional difference during this period (as discussed in Section 2), one could hardly expect a VAR such as this to account for all bank panics predictors, and as such, I find increases in the likelihood of panics occurring from an aggregate demand shocks being 3.4% (and slightly higher or lower depending on the alternative specifications below) but not substantially more. Nonetheless we do see that unanticipated deflation can make panics more likely, but only when not offset by higher-than-expected real output. In other words, negative aggregate demand shocks increase the likelihood of bank panics, though positive aggregate supply shocks do not, even though both shocks lead to a lower-than-expected price level.

One potential concern with these findings is, given that a negative aggregate demand shock in this specification requires a reduction in real output, perhaps it is the reduction in real output causing the increased probability in bank panics. Yet, as shown by Jalil (2015), testing for both output shocks and price shocks in an unrestricted VAR with the same endogenous variables finds that neither, *on their own*, has any significant impact on the likelihood of bank panics. Furthermore, given that impulse response functions derived from a VAR model are symmetrical, it follows that if falling output on its own could be driving the effect, we would expect to see a result significantly different from zero in Figure 3, which we do not. This suggests that declines in real output only have the ability to increase the probability of bank panic if prices do not rise sufficiently offset the decline.

In Figures 4 and 5, I present the same impulse response functions as before, but this time I do not seasonally adjust the Long Construction to more closely align with the set-up of Jalil (2015). This is to help verify that what is driving our differing results is not from differences in VAR inputs, but from the way in which I've isolated the shocks. Here we see the same general result but with a stronger effect on impact of an aggregate demand shock on panics of 8.4%. This second specification is using the same inputs for output and prices as Jalil,[25] and therefore we can attribute the fact that he found no effect to a price shock on panics to the fact that he did not decompose price shocks by their co-movement with output as done here.

A complete understanding as to why the seasonal adjustment of output data lowers the effect of procyclical shocks on bank panics is beyond the scope of this paper. However, recall that for the debt-deflation theory to lead to bank panics, the deflation need to be

[25] My dummy variable is still slightly different than in Jalil's VAR, where he does not include nonmajor panics. However, as I show in Appendix A.2, excluding nonmajor panics does not significantly change the results.

Figure 4: Response to Negative AD Shock (Not Seasonally Adjusted)

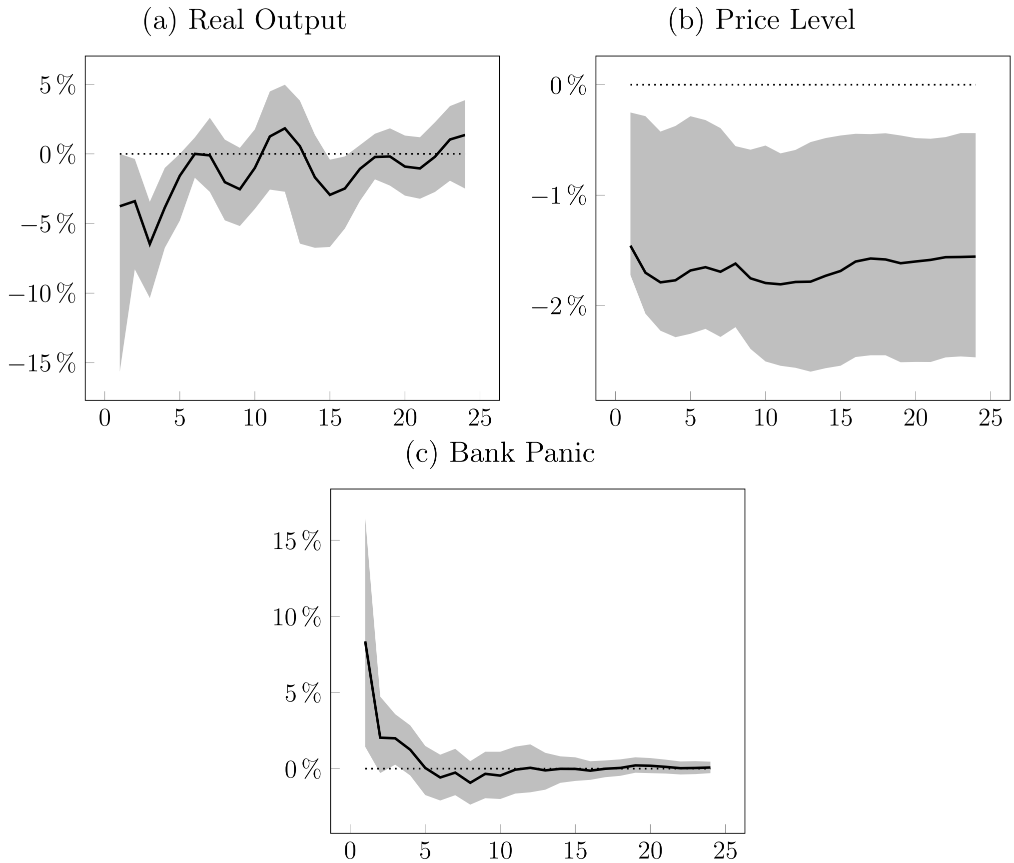
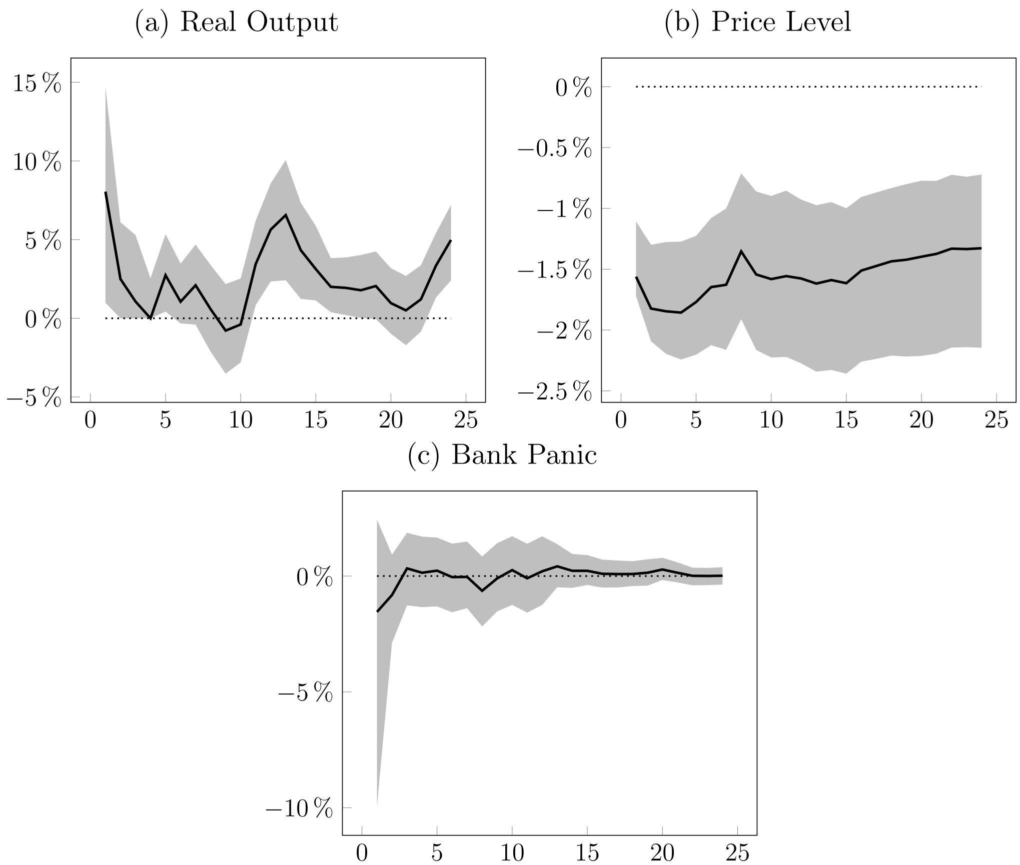


Figure 5: Response to Positive AS Shock (Not Seasonally Adjusted)



unanticipated. It is possible that during this period—perhaps do to lack of real-time data availability—seasonal fluctuations were not as expected as we believe them to be today. If that were the case, then by seasonally adjusting the data, it is possible I removed some of the underlying uncertainty from the data.

One conceivable way to better capture movements in prices and output which are truly unanticipated is to remove a trend component from the data before estimating the VAR. While anticipating seasonal movements in the 1800's may have been difficult, it is plausible people could still have extracted long run trends. If borrower and lenders agreed on the long run tend of prices when agreeing upon debt contacts then this could be priced into their deal. For example, as touched upon in Section 3, for about 30 years prior to 1896 mild deflation was the norm. It therefore seems reasonable to believe that much of this deflation was anticipated.

To account for this, I present the impulse response functions in Figures 6 and 7, where a time-trend has initially been removed from the data. To remove a trend component from both output and prices, I use the Hamilton Filter.[26] Removing the trend component again turns out the same basic result. Again we observe that negative aggregate demand shocks (*i.e.,* unanticipated deflation, not offset by output increases), significantly impacts bank panics. Though again, in this case, the magnitude is lower then when output and price data are unadjusted. In this case we also see a small significant effect on bank panics on impact from a positive supply shock, which could suggest that in this case output volatility is driving more of the result in this specific specification. However, even though it is significant on impact (in that the 16th and 84th percentile bands do not straddle zero), the magnitude is small and quickly returns to zero.

Despite the correlation of the Long Construction Index with the Miron-Romer Index,

[26]These results are robust to using other filters such as Hodrick-Prescott or Baxter-King. However, here I present the results with the Hamilton filter because of the arguments in Hamilton (2018).

Figure 6: Response to Negative AD Shock (De-trended with Hamilton Filter)

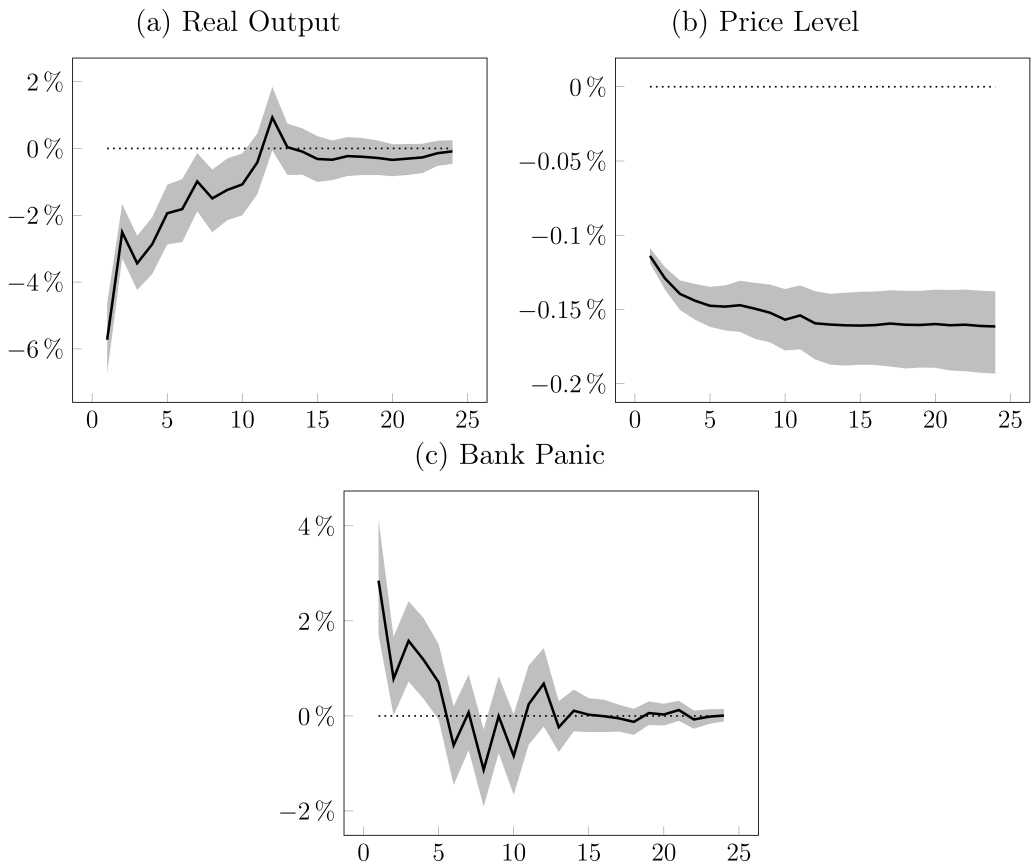
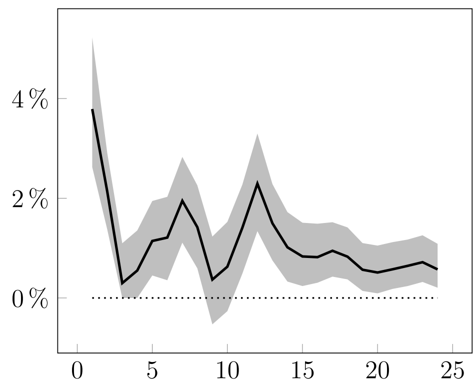
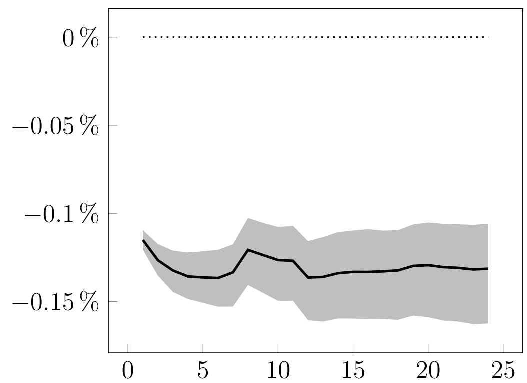


Figure 7: Response to Positive AS Shock (De-trended with Hamilton Filter)

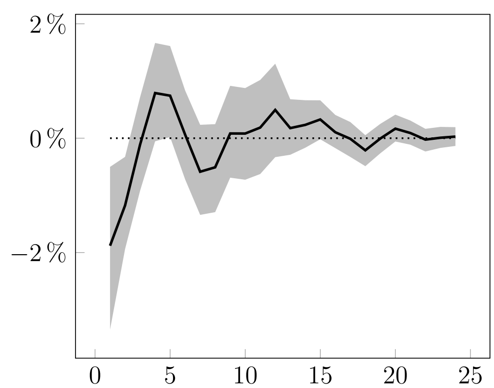
(a) Real Output



(b) Price Level



(c) Bank Panic



concerns may still arise regarding the appropriateness of a construction index as a proxy for total real output. To account for this, I estimate the same sign restricted VAR, this time using the quarterly real GNP and GNP deflator estimates from Balke & Gordon (1989) as my measure of output and prices. This truncates the beginning period of study somewhat to Q1 of 1875 due to data availability, though much less so than if the Miron-Romer data was used directly.[27] Looking back to Table 1, this means the major bank panic of 1873 can no loner be included in the panic dummy. The dummy variable for panics must be further modified to match the quarterly frequency of output and prices in this case. Therefore, under this specification, b now includes a "1" in any quarter where a bank panic began, and a zero else.[28]

Figures 8 and 9 present the impulse responses for 16 periods, or four years out. Like with the monthly data, only aggregate demand shocks significantly increase the likelihood of a bank panic beginning. With this specification, panics are 5% more likely to begin on impact of an aggregate demand shock, and remains significantly above zero for one quarter. While information is clearly lost when using a lower frequency and a later starting date, what is gained is the ability to look at total nominal income. Having actual real output and price level estimates, instead of rough proxies, allows me the ability to see to what extent output changes are offsetting price level changes during aggregate supply shocks. By definition, NGDP must be lower-than-expected on impact from an aggregate demand shock, what is interesting however, is analysing NGDP during a supply shock. Looking to Figure 9 we

[27]While the Balke & Gordon (1989) annual series begins in 1869, the quarterly data does not begin until 1875. In theory the VAR could be run at the annual level, however, I argue that is likely too low of a frequency to reasonably capture the effects under study. This is evidenced by the fact that the vast majority of bank panics in this period lasted only one month.

[28]Note that under the monthly specifications, "1"s were given for any month which a bank panic was occurring, not only much in which a bank panic began. However, this leads to problems with aggregation to the quarterly data. For example, a hypothetical shock lasting January-March (three months long) would only count as a single "1" when aggregated to quarterly. However, a shock lasting through March and April of another year (only two months long) would be counted as two "1"s (in Q1 and Q2) once aggregating to quarterly. Using only months in which a panic begins avoids this problem.

Figure 8: Response to Negative AD Shock (Quarterly Balke-Gordon Data)

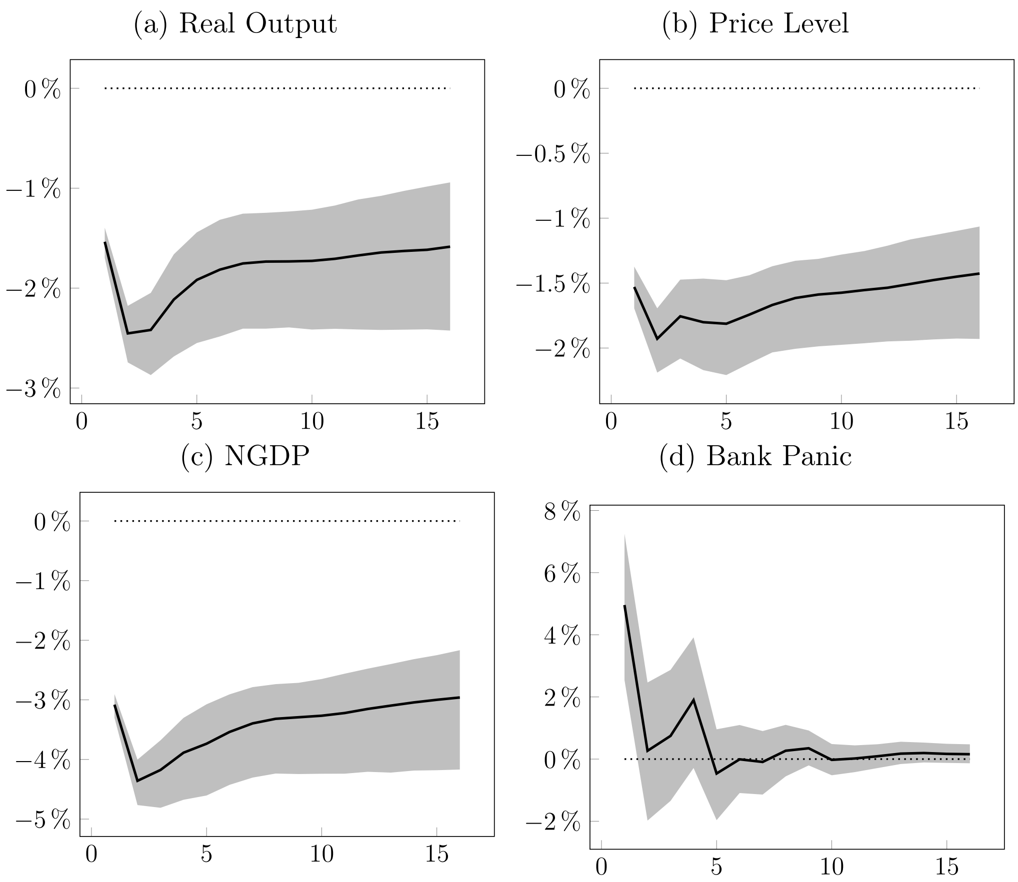
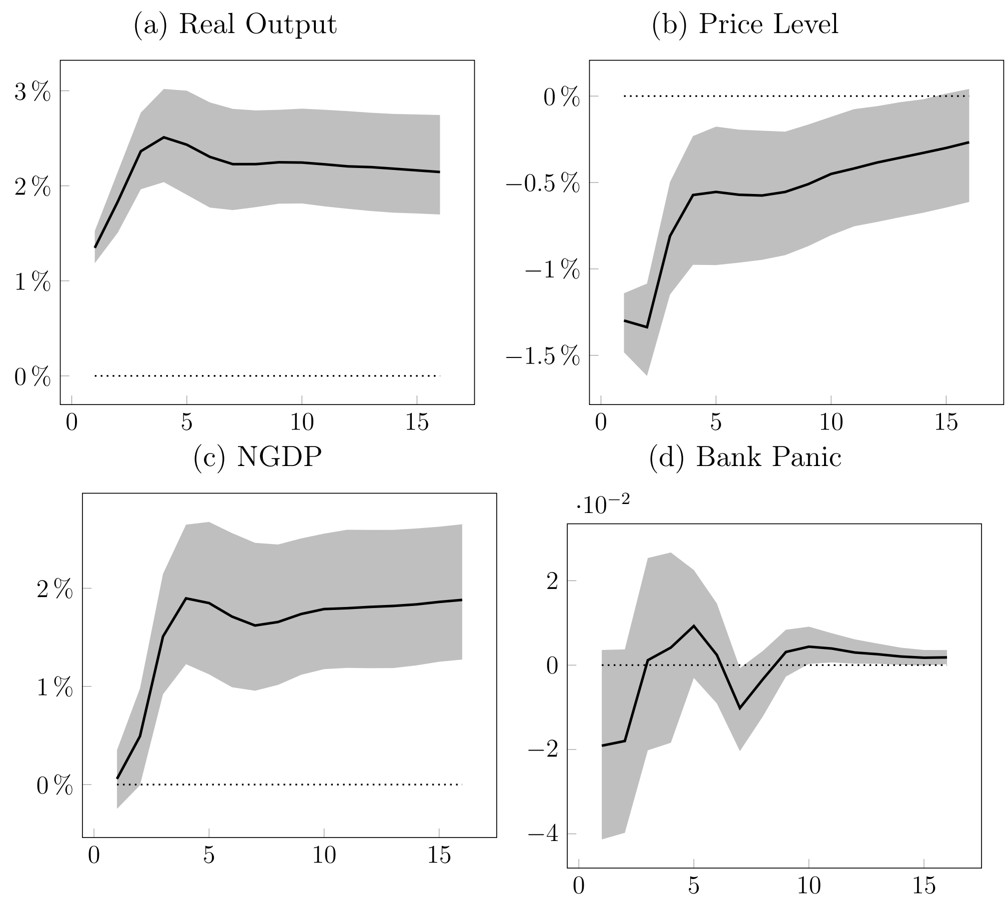


Figure 9: Response to Positive AS Shock (Quarterly Balke-Gordon Data)



see that output movements more than offset the negative price shock, causing NGDP to be positive on impact. Thus, while the result is not significant, the fact that in this case the median response on panics is negative, is likely driven by the higher-than-expected nominal income.

The combined results presented in Figures 2 - 9 show that, to the extent debt-deflation can lead to bank panics, it is only when not offset by higher-than-expected real output, or in other words only when there is a collapse in aggregate demand. Theory put forth by Selgin (1997), Sumner (2012), Koenig (2013), Sheedy (2014), Azariadis et al. (2019), Bullard & DiCecio (2019); and Bullard & Singh (2020) suggests that when unanticipated deflation is caused by positive productivity shocks—or any positive supply shock which reduces the marginal costs of the average firm without cutting wages—the resulting increase to real incomes will offset any increase to the real value of debt. In that case, debt-to-income ratios are able to remain stable despite prices not being stable and therefore there is no increase in the likelihood of bank panics despite the unexpected deflation. This section has presented empirical results to bolster this theory.

6 Conclusion

Understanding what causes banking panics is important because when bank panics occur their effects tend to stretch far beyond the financial sector, impacting many facets of the real economy including causing decreases in employment and output (Friedman & Schwartz 1963, Bernanke 1983, Grossman 1993, Jalil 2015). This paper contributes to our understanding by offering evidence that challenges the blanket assertion that all unanticipated deflation can increase the likelihood of banking panics. I chose the U.S. National Banking era—marked by frequent deflation and bank panics—as my period period for this analysis.

I employed a structural VAR model with sign restrictions to discern two unique deflationary shocks: a negative aggregate demand shock, where output and prices move together on impact, and a positive aggregate supply shock, where output prices move in the opposite direction on impact. Depending on the specifications applied, the impulse response functions show an increase in the likelihood of a bank panic occurring by 3.4 to 8.4% in response to a negative aggregate demand shock, with no significant effect to a positive aggrage supply shock. This result suggests that the connection between debt-deflation and banking panics is not a universal phenomenon, but rather occurs when unanticipated deflation is not offset by rises in real output.

Bank panics should be avoided if possible, due to the real harm they cause when they occur. If we know their cause, then perhaps policy can be implemented which aids in preventing them. Much has changed between 1913 and today, limiting our ability to draw conclusions from the past with complete confidence. Nonetheless, these results provide evidence in favour of theories such as Koenig (2013), Sheedy (2014), Azariadis et al. (2019), Bullard & DiCecio (2019); and Bullard & Singh (2020) arguing that nominal income stability—being accommodative to aggregate supply driven deflation—is the more important metric than price stability when attempting to avoid unexpected defaults and delinquencies and even bank panics. These results also seem complement other works such as Hayek (1935, 1960); Selgin (1988, 1997); Bordo & Redish (2004); White (2006); Beckworth (2007); Beckworth (2014); and Beckworth (2019) suggesting that only when general changes to the price level stem from unstable aggregate demand, is there an ability to cause real economic harm, whereas general price level changes from the supply side are seen as benign or even beneficial as the price system at work.

Combining the conclusions of this body of literature with my findings here seems to suggest that absolute price stability or monetary policies intended to prevent any unexpected deflation might be excessive, at least regarding ensuring the smooth functioning of financial intermediation. Indeed, these results not only support the idea that certain unanticipated deflation–specifically deflation stemming from unexpected productivity gains leading to increased real output and higher-than-expected real incomes, *i.e.,* a positive supply shock—may be perfectly compatible with financial deepening. Other works from Hayek (1928), Selgin (1997), Sumner (2012), Beckworth (2014), Koenig (2013); and Sheedy (2014) have demonstrated that monetary policy aimed at counteracting this type of supply-driven deflation can introduce new problems of it own.

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A Appendix

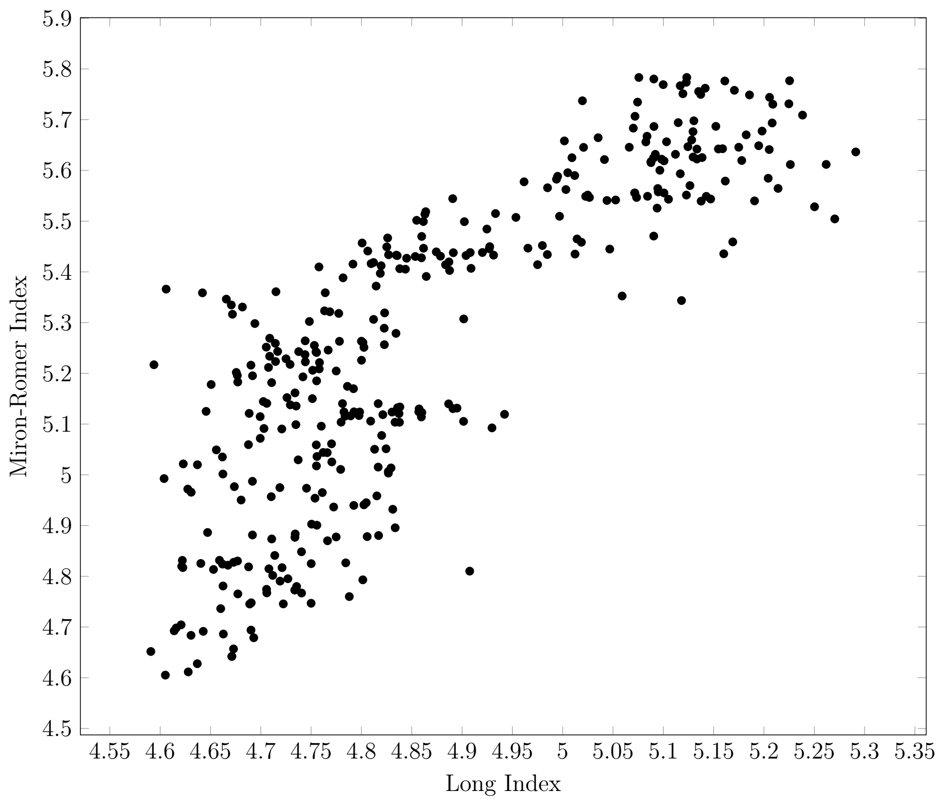
A.1 Correlation between the Long Index and the Miron-Romer Index

As a further check for the appropriateness of using the Long Index as a proxy for output, I check its correlation against the Adjusted Miron-Romer Index of Industrial Production (Miron & Romer 1990), which is a well-know broad index giving a more accurate sense of monthly real economic activity (see Stock & Watson (1999) and Gordon (2016) for examples of its implementation as a measure of real output).[29] The Miron-Romer Index does not begin until July 1884, however, omitting a large portion of the National Banking era, including 1973 when a major panic took place (see Table 1). While I cannot therefore use it as a measure of output for the period of study here, if it is well correlated with the Long Index for months between July 1884 and December 1913, then we should be reasonably assured

[29]The Miron-Romer Index includes the following industries: coal, petroleum, sugar, cattle, pigs, coke (another form of coal), flour, wool, coffee, tin, rubber, and silk (Miron & Romer 1990).

that the Long Index is capturing real economic activity as well. Indeed, looking at the covariance matrix between the two series, I find a value of 0.852, suggesting that the series' are well correlated. Furthermore, looking to Figure 10, suppressing the constant term and regressing the Long series against the Adjusted Miron-Romer series gives a beta estimate of 1.06 and an R^2 of 0.998.

Figure 10: Correlation between Construction Activity and Production Jul 1884- Dec 1913 (Jul 1884 = 100, log scale)



Source: Long (1940) for construction, and Miron & Romer (1990) for production.

A.2 A VAR Including Major Panics Only

As mentioned a number of times throughout this paper, much of the VAR setup was explicitly meant to mimic the setup found in Jalil (2015). This was done to better isolate how our contrasting results can be distinguished – to provide evidence that our differing results are from the way in which I have uniquely unidentified my shocks, not due to different input into the VAR. However, throughout 5, my dummy variable for panics differed from Jalil's in that I included nonmajor panics. This was on the assumption that the varying levels of fragility in each States banking system may have lead deflation to lead to panic in some regions without spreading to others. It was also done to add more variation to the dummy variable, as there are only three major panics in a series of 552 months.

However, here I present the results with only major panics included in the dummy as well as the unadjusted Long Index and the USA Annalist Wholesale Price Index.

As can be seen, this does not change the main findings of the paper, and given the same inputs, this further confirms that my differing results are driven by the decomposition of the two deflations.

A.3 A VAR Including Alternative Price Data

While Jalil also uses the USA Annalist Wholesale Price Index, as do I, other monthly price series are available for the National Banking era. The U.S. Index of the General Price Level from *NBER Macrohistory Database* (2022), being made up of broader basket of goods than only wholesale prices,[30] arguably gives an even better measure of deflation in the traditional sense of "a fall in the general price level". Figure A.3 plots the U.S. Index of the General Price Level and the the USA Annalist Wholesale Price Index. When checking their correlation,

[30] The U.S. Index of the General Price Level is composed of industrial prices, prices of non-agricultural good, farm prices, wholesale prices, retail food prices in 51 cities, rents in 32 cities, clothing, fuel, furnishings, freight and transportation costs, realty values, securities, bonds and stocks, equipment and machinery, hardware prices, automobiles prices, and wages (*NBER Macrohistory Database* 2022).

Figure 11: Response to Negative AD Shock (without nonmajor panics)

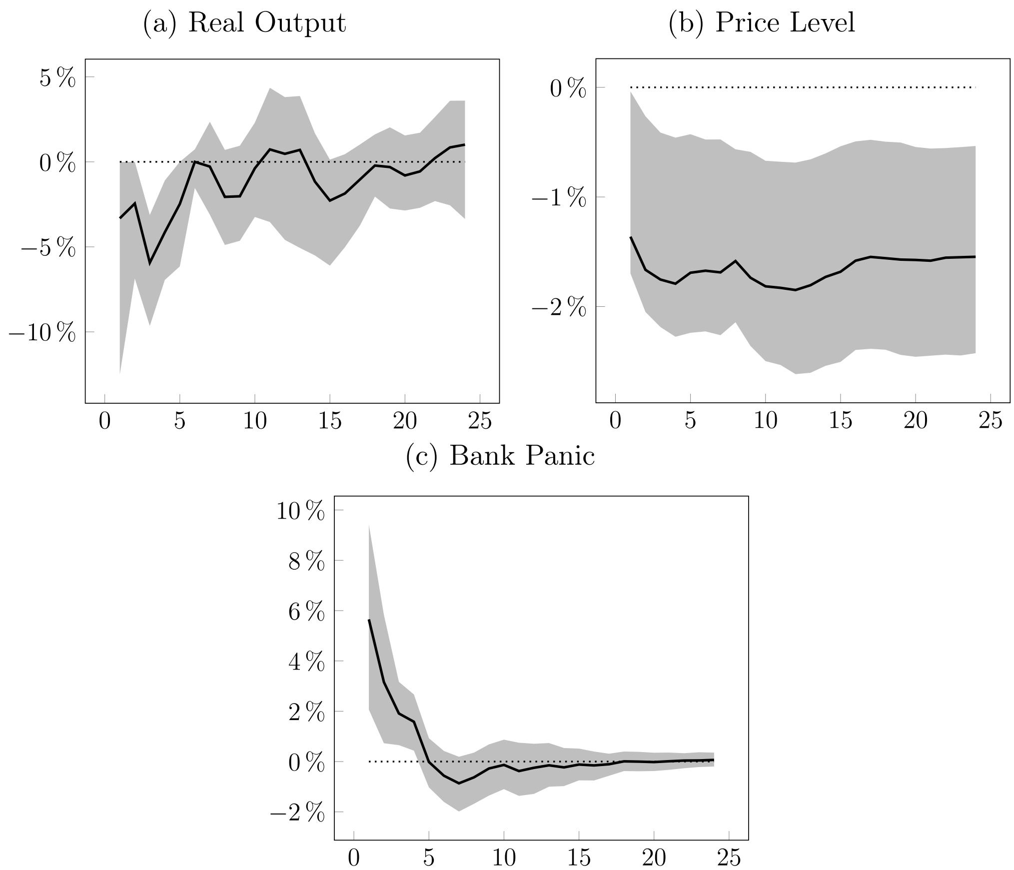
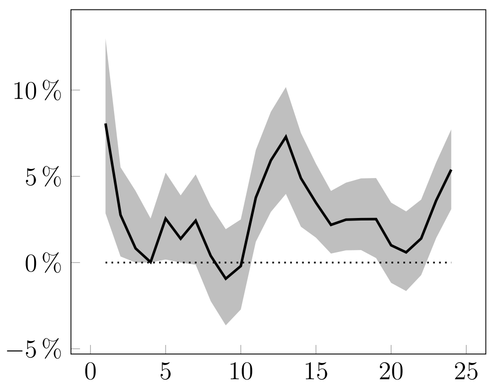
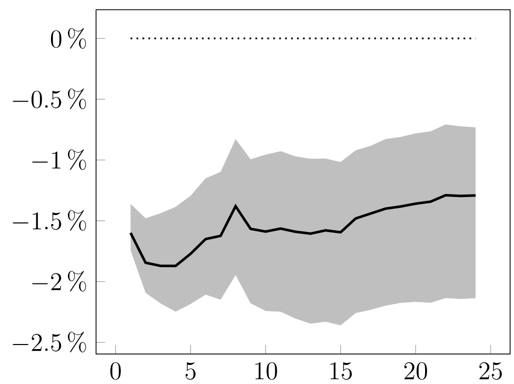


Figure 12: Response to Positive AS Shock (without nonmajor panics)

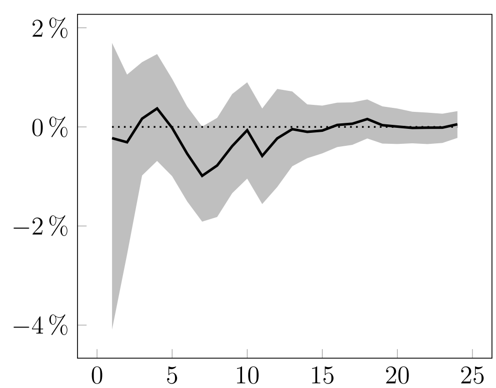
(a) Real Output



(b) Price Level

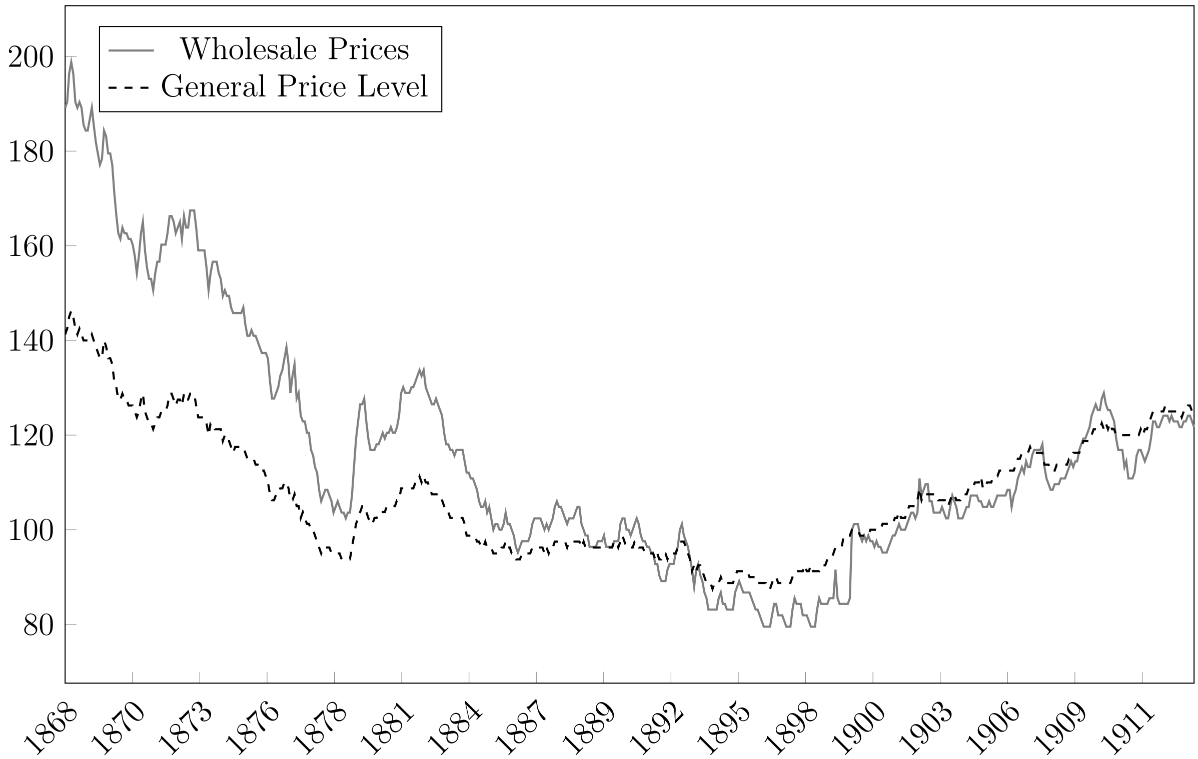


(c) Bank Panic



I get a value of 0.80. Regressing one against the other I get a R^2 of 0.98 and a β of 1.13. Suggesting that these two price series are well correlated, and so using either should obtain similar results.

Figure 13: Alternative Data on Prices



Source: Jalil (2015) for Wholesale Prices, and NBER Macrohistory Database *(2022) for the General Price Level*

Note: January 1900 = 100 for both series.

The impulse responses obtained from using the general price level instead of wholesale prices are presented in 15 and 14. Yet again, we see the same key result presented, where only demand shocks increases the likelihood of panics occurring.

Figure 14: Response to Negative AD Shock (with General Price Level)

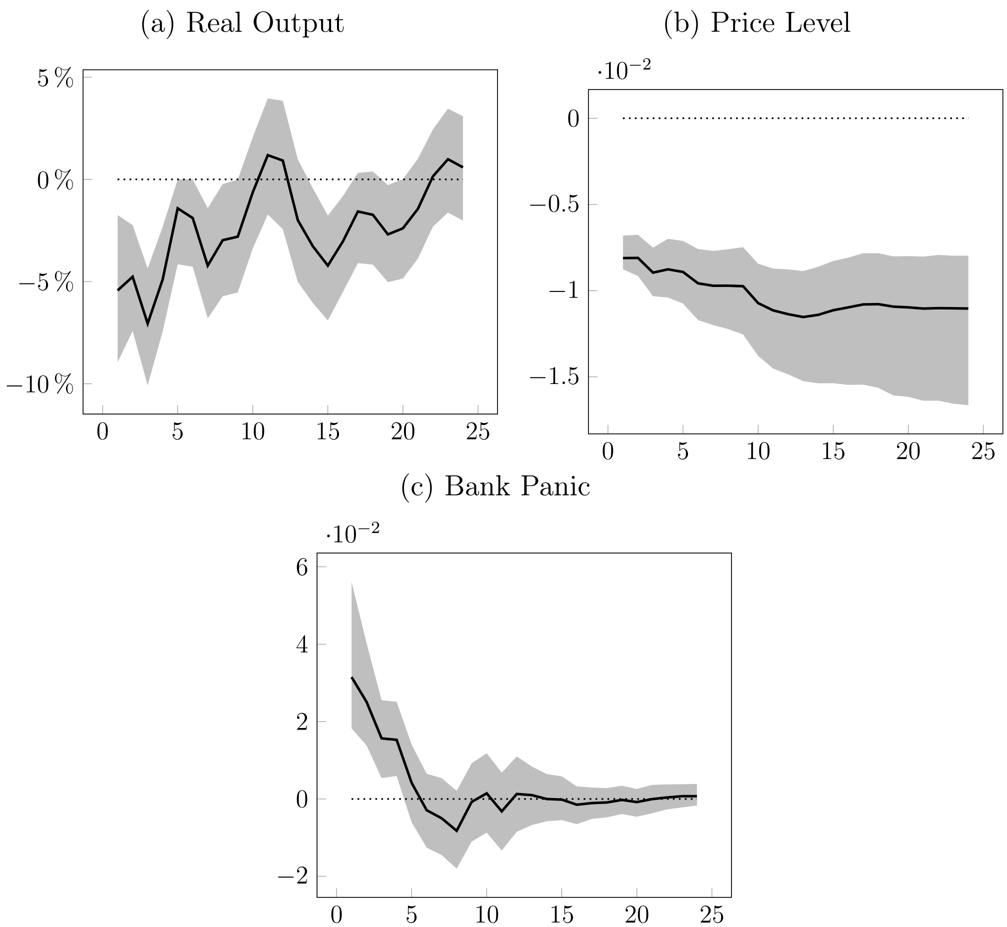
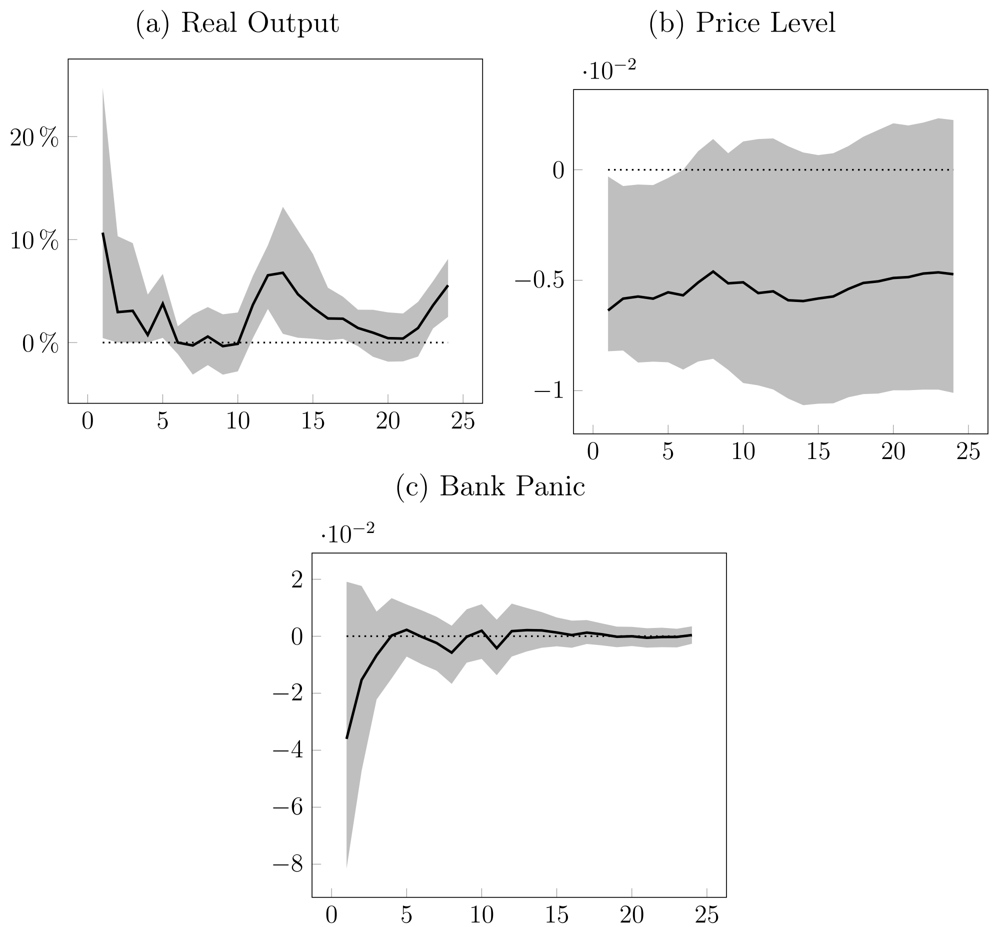


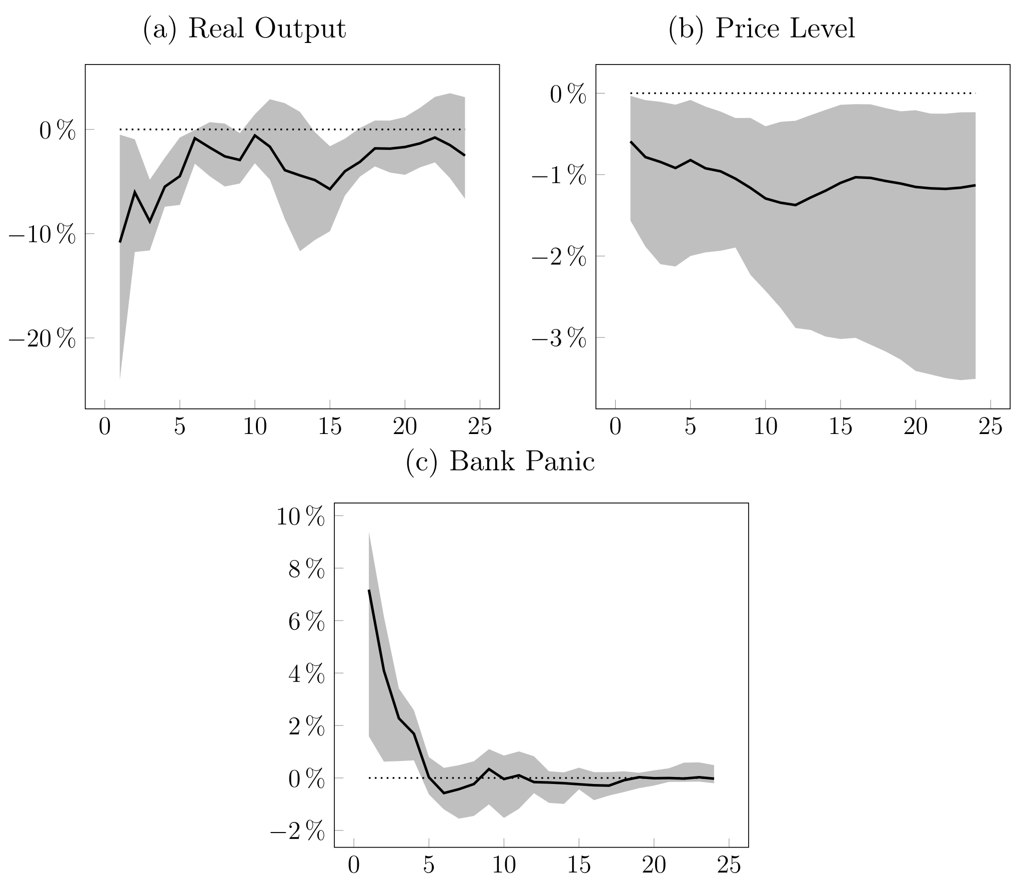
Figure 15: Response to Positive AS Shock (with General Price Level)



A.4 A VAR using the Rubio-Ramirez et al. (2010) Rejection Method

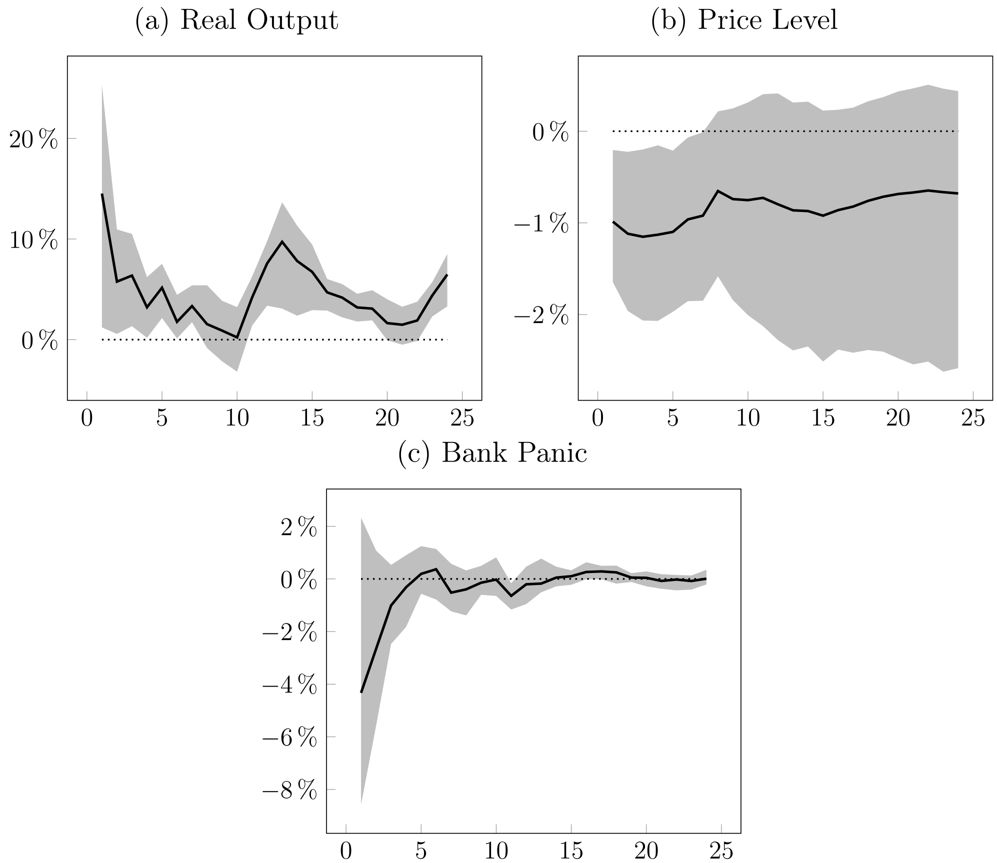
In order to obtain 1000 sign-conforming draws for the results in Section 5 I implemented the use of a penalty function from Uhlig (2005), and detailed in Section 4. However, Rubio-Ramirez et al. (2010) have proposed alternative method. This alternative rejection method is comprised of an algorithm for accepting C matrices whose impulse responses match the directions of sign restrictions while rejecting those which to not. Because the Rubio-Ramirez et al. (2010) method will yield a different set of 1000 results, it's possible that the overall effect could be different.

Figure 16: Response to Negative AD Shock (Using RWZ Rejection Method)



Looking to Figures 17 and 16, however, while some of the median magnitudes are altered, the key results remain. Even when using this rejection method in favour of a penalty function

Figure 17: Response to Positive AS Shock (Using RWZ Rejection Method)



(a) Real Output

(b) Price Level

(c) Bank Panic

we still see a significant increase in the likelihood of bank panics during a negative aggregate demand shock but not during a positive aggregate demand shock. In fact, the median response of panics to a demand shock has only increased from the original specification, rising from 3.4% to 7.2%.