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Financial Crisis and the Great Depression: A Regime Switching Approach

I explore the timing of and effects of the U.S. financial crisis of the 1930s in a regime switching framework. Estimated conditional probabilities over the state of the financial sector suggest that a prolonged period of crisis begins not with the 1929 stock market crash, but with the first banking panic of October 1930. These probabilities also suggest that the crisis persists until the introduction of federal deposit insurance in early 1934. Consistent with the view that this financial crisis had real effects, these conditional probabilities contain additional explanatory power for output fluctuations. This is in addition to that provided by the money stock.

RECENT EVENTS IN SOUTHEAST ASIA have focused much attention on financial crises. Despite this attention, there is no consensus view as to what causes financial crises, what are the appropriate policies for ending a crisis, and what are the real effects of such crises. History provides several examples of financial crisis from which one can draw inference.

Perhaps the most notable and most studied of these occurred in the United States during the 1930s. However, even in this case, there are a number of possible causes and solutions. Potential causes include a widespread contagion of fear among depositors after the failure of a few large banks and the stock market crash of 1929. Candidate solutions include the bank holiday and Roosevelt's fireside chats, the suspension of the gold standard, and the introduction of federal deposit insurance. Finally, it has been argued that this financial crisis had effects on the real side of the economy, both through a collapse in the money supply (Friedman and Schwartz 1963) and through a collapse in credit intermediation (Bernanke 1983).

The purpose of this paper is twofold; to make inferences about the timing of the 1930s' U.S. financial crisis and to test whether or not it had real-side effects. In order to do this I assume that, at any point in time, the financial sector is in one of two regimes, crisis or calm. Using a bivariate version of Hamilton's (1989) Markov switching model, I then estimate a conditional probability of financial crisis for each

The author thanks, without implicating, Anji Redish, Jim Nason, Herb Emery, Tim Hatton, Joe Mason, Ron Shearer, and an anonymous referee for their comments and advice. Financial support from ESRC research grant R000235524 is gratefully acknowledged.

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Journal of Money, Credit, and Banking, Vol. 34, No. 1 (February 2002)

date in my sample. To test for real effects I include this time series of estimated probabilities in a reduced-form equation for output growth.

This gives me some insight into the areas suggested above. First, by estimating a time series of conditional probabilities over the state of the financial sector I can draw inferences about the timing of the financial crisis. This allows me to see which events coincide with the onset of financial crisis and whether any policy reforms coincide with its end. My evidence suggests that the financial sector did not move into a prolonged state of crisis until the first wave of banking panics in October 1930. This is consistent with Mishkin's (1991) view that a stock market crash, in this case, the 1929 crash, does not necessarily imply a financial crisis. Perhaps the more interesting result regarding the timing of the financial crisis pertains to its end. The traditional view has the crisis ending in the Spring of 1933 with the bank holiday, Roosevelt's fireside chats, and the abandonment of the gold standard.[1] The results in this paper, however, suggest the financial crisis persists until the introduction of federal deposit insurance in early 1934.

Second, by using the estimated probabilities over the state of the financial sector as explanatory variables in output equations, I am able to test whether the financial crisis had significant effects on the real side of the economy. I find that these probabilities do have additional explanatory power over an autoregressive process for manufacturing production. Consistent with Bernanke's view that the effects of the financial crisis were not solely through the monetary channel, I find that this explanatory power remains when the growth rate of M2 is added to the equation. Recently Cooper and Ejarque (1995), and Cooper and Corbae (1999) have suggested that the interwar financial sector can be characterized as shifting between two states that are indexed by agents' confidence in the intermediation process. The evidence of regime switches I find in this paper is consistent with such an approach. Also consistent with this approach's ability to further explain interwar output fluctuations, I find that the probabilities of financial crisis retain their explanatory power once linear measures of financial crisis are added to the model.

In the next section I outline the channels via which it has been argued that the disruptions to the financial sector had real effects. In this section I also introduce the two measures of financial crisis that I use in the empirical section and provide a brief review of policy responses to the financial crisis in the early 1930s. Section 2 uses the Markov switching model to explore the timing of the crisis and section 3 tests for real effects of the crisis. Section 4 concludes.

1. FINANCIAL CRISIS AND FINANCIAL REFORM

It has long been argued that disruptions to the financial sector were an important source of fluctuations in real activity during the Great Depression. In this section I

1. See, for example, Burns (1974) or Kennedy (1973) for an emphasis on domestic reforms, or Wigmore (1987) for emphasis on the decision to abandon the gold standard.

outline the monetary and nonmonetary channels via which it has been argued that the financial crisis contributed to the Great Depression, motivate two measures of financial crisis and discuss policy responses to the crisis.

1.1 Effects and Measures of Financial Crisis

Friedman and Schwartz argue that the bank failures of the 1930s had an adverse effect on income during the Great Depression through two channels: a negative wealth effect for bank shareholders and, much more importantly, through a contraction in the money supply.[2] According to Friedman and Schwartz the bank failures of the 1930s caused households to substitute away from deposits and into other assets such as currency. Through the money multiplier this lead to a decline in the money stock. With short-run non-neutralities of money, this may have caused some of the decline in real income.

Friedman and Schwartz's argument suggests that the deposit-currency ratio would have declined during the financial crisis. Figure 1a shows the monthly growth rate of the deposit-currency ratio between 1919 and 1941. The early 1930s stands out as a period in which there were large negative growth rates in the deposit-currency ratio.[3] (For details of this and all other data used in this paper, see Appendix 1.)

Bernanke proposes a third channel through which the financial crisis had an adverse effect on income: the collapse of credit intermediation. Banks typically have mostly illiquid assets (loans) and demandable liabilities (deposits). During a bank run, solvent banks can be caught in a liquidity crisis and be forced to suspend payments. Bernanke argues that as banks began to fail, nonfailing banks began to fear a run. Hence they increased their reserve-deposit ratios and substituted into safer and more liquid assets. This led to a reduction in their role of credit intermediation and so a rise in the cost of credit intermediation. Consequently the borrower's rate rose and the number of loans fell, leading to a decline in real activity.

Bernanke's argument implies that during a financial crisis the cost of credit intermediation will be higher. Unfortunately, as he points out, there is no direct measure of the cost of credit intermediation available.[4] Here I follow Bernanke (1983) and Mishkin (1991), and use a yield spread as a proxy. This spread is the difference between the yield on corporate bonds rated Baa by Moody's and the yield on long-term government bonds. Although this is not a perfect measure of the cost of credit intermediation, it does represent the different costs of funds to two different classes of

2. Temin (1976) finds little evidence to support the hypothesis that there was a negative wealth effect following the stock market crash of 1929. He attributes this to the fact that stocks were only a small proportion of wealth. This also suggests a minor effect on aggregate income via a negative wealth effect following the bank failures.

3. I also experiment with the real liabilities of failed banks, the growth rate of nominal deposits, and the growth rate of real deposits in place of the growth rate of the currency-deposit ratio. The results of the paper do not change. Full details of these and other results mentioned, but not reported, are available from the author on request.

4. If banks were making only the safest and highest-quality loans, the reported rates for commercial loans do not reflect the shadow cost of bank funds to a representative borrower.

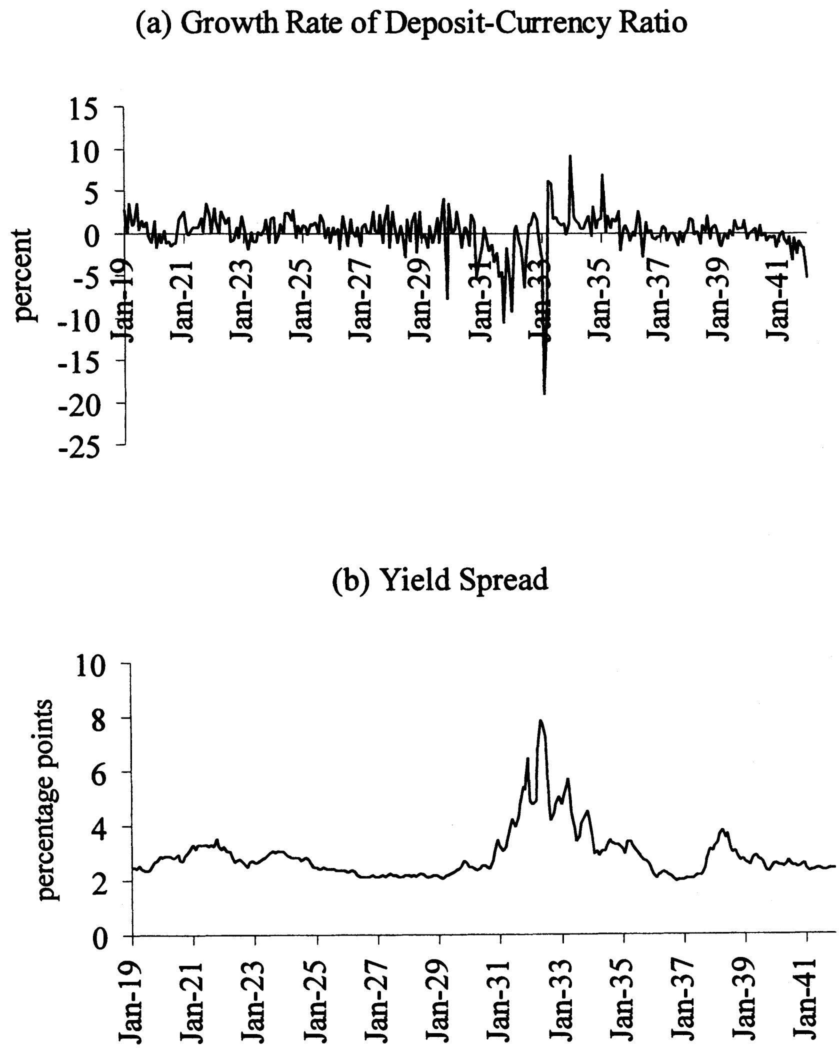

FIG. 1. Growth Rate of Deposit-Currency Ratio and Yield Spread, 1919–1941

borrowers.[5] Figure 1b shows that the spread between Baa rated bonds and long-term government bonds was higher during the Great Depression than during the rest of the interwar period. Also consistent with a financial crisis is the observation that this increase in the spread comes from an increase in the yield on the risky asset rather than a fall in the yield on the risk-free asset. In January 1930 the yield on Baa rated bonds was 5.29 percent. It reached a peak at 11.63 percent in May 1932, which is also when the spread peaks. In January 1930 the return on long-term government bonds was 3.43 percent. In May 1932 it was relatively unchanged at 3.76 percent.

Given the links between financial conditions and the real economy, I seek to further understand the Great Depression by focusing on financial conditions in the U.S. economy in the 1930s. I suggest that in the interwar period the U.S. financial system can be characterized by two states: first, the financial crisis state, in which the fear of bank failure was high and all but the safest class of borrower found it expensive to obtain credit, as reflected in a declining deposit-currency ratio and a higher cost of credit intermediation. A second state, the financial calm state, has a low incidence of bank failure and more equal costs of credit across borrowers. In this state the deposit-currency ratio was stable and the cost of credit intermediation was lower.

1.2 Policy Responses to the Crisis

This section provides a brief overview of the policy responses to the financial crisis.[6] During 1932 there were two major policy initiatives aimed at alleviating the financial crisis, although neither appears to have had the desired effect. The first was the introduction of the Reconstruction Finance Corporation (RFC) in January 1932 to provide loans to illiquid banks. However, Mason (1996) argues that by overcollateralizing these loans, the RFC actually created a liquidity problem for the very banks it was trying to help. The second response was the Federal Reserve's open market purchases, which began in April 1932, but were abandoned in July of that year. Epstien and Ferguson (1984) argue that this program was abandoned due to pressure from member banks. As banks were substituting away from loans and into short-term government securities the Fed's open market purchases had an adverse effect on their profitability.[7] As a result they became increasingly opposed to the program.

At midnight on March 6th, 1933, the newly inaugurated President Roosevelt declared that there would be a bank holiday from the 6th to the 9th of March. On the 9th of March, Congress passed the Emergency Banking Act (EBA) which was the first of many banking and monetary reforms contained in the New Deal. The EBA gave the RFC power to invest equity in banks without taking collateral, thus solving the problem Mason discusses. The EBA also facilitated the reopening of national banks. Roosevelt promised the public that only "sound" banks would be granted li-

5. I also use the yield on Moody's highest rated bonds (Aaa) in the place of the yield on long-term government bonds. Again, my results are unaffected.

6. For a much fuller chronology of the bank failures, policy responses, and other important events during the Great Depression, see Friedman and Schwartz (1963, Chapter 7) or Wicker (1996).

7. The yield for three- to six-month treasury notes and certificates fell from 2.25 percent in March 1932 to 0.22 percent in July, *Banking and Monetary Statistics* (1943, Table 122).

censes to reopen. These reforms, the program for reopening the banks, and Roosevelt's "fireside chats" were intended to stabilize the financial system. The traditional view emphasizes the success of these measures in restoring stability to the financial sector. See, for example, Kennedy (1973) or Burns who writes:

> The Roosevelt forces moved quickly, firmly, and with courageous optimism to reassure the people. When the banks reopened, a surge of confidence swept the country and carried in its wake the passage of the Banking Act of 1933. (1974, p. 181)

Friedman and Schwartz argue that by restoring confidence in the monetary and economic system the EBA contributed to recovery from the depression. However, they also argue that the introduction of the Federal Deposit Insurance Corporation (FDIC) was the structural change that did the most to restore stability. This occurred in January 1934. By July 1934 almost 14,000 of 15,348 commercial banks were covered.[8] Friedman and Schwartz argue that the number of bank failures was greatly reduced in 1934 for two reasons. First, small depositors knew that if their bank failed they would be reimbursed; therefore, bank runs did not spread from one bank to another. Second, "bad" banks were not allowed to fail; instead they were merged with "good" banks or re-organized under new management with the FDIC assuming any losses.

An alternative view is offered by Wigmore (1987). He questions whether the banking reforms of the EBA and the introduction of the FDIC Act were sufficient to explain the subsequent calm in the financial sector. He points out that the FDIC only covered accounts up to \$2500, left \$25bn in large accounts uncovered (two-thirds of total deposits in insured banks) and did not become effective until 1934. The ability of the RFC to provide capital is also seen as relatively unimportant by Wigmore, as only \$15m of authorizations had been made by the end of March 1933. On the other hand, Mason points out that there was a substantial increase in RFC authorizations in late 1933 and early 1934 as banks were recapitalized in preparation for joining the FDIC. Wigmore goes on to argue that much of the calm can be attributed to the suspension of the gold standard in March 1933.

2. TIMING OF THE FINANCIAL CRISIS

Using a Markov-switching model I now seek to identify periods of financial crisis and financial calm between 1919 and 1941. This allows me to draw insight as to when the financial crisis starts and when it ends, and to see which events, if any, coincide with these changes in regime. To do this I estimate a conditional probability of financial crisis for each month between May 1919 and December 1941 using a bivariate Markov-switching model. This probability is conditional on the two measures of financial crisis discussed in the previous section. I then discuss the timing of the crisis implied by these probabilities.

The model I present below is a variant of Hamilton's (1989) Markov-switching

8. Friedman and Schwartz (1963, pp. 436–37)

model. Two stationary time series follow a common regime, S_t.[9] Assume that the vector, $\mathbf{x_t}$ consisting of the growth rate of the deposit-currency ratio ($x_{1,t}$) and the yield spread ($x_{2,t}$) follows the stochastic process:

$$\begin{bmatrix} x_{1,t} - \mu_1(S_t) \\ x_{2,t} - \mu_2(S_t) \end{bmatrix} = \sum_{j=1}^{n} \begin{bmatrix} \rho_{1,j} & 0 \\ 0 & \rho_{2,j} \end{bmatrix} \begin{bmatrix} x_{1,t-j} - \mu_1(S_{t-j}) \\ x_{2,t-j} - \mu_2(S_{t-j}) \end{bmatrix} + \begin{bmatrix} \sigma_1(S_t)\nu_{1,t} \\ \sigma_2(S_t)\nu_{2,t} \end{bmatrix}. \qquad (1)$$

Here μ_1 refers to the mean of the first series, the growth rate of the deposit-currency ratio, and μ_2 refers to the mean of the second series, the yield spread. The parameters $\rho_{1,j}$ and $\rho_{2,j}$ (where $j = 1,2,\dots,n$) are the autoregressive parameters for the two series respectively, and $n = 4$ is the number of lagged dependent variables.[10] The error terms, $\nu_{i,t} \sim i.i.d.N(0,1)$, where $i = 1,2$, and $\sigma_1(S_t)$ and $\sigma_2(S_t)$ are their standard deviations.

This model differs from the standard autoregressive model as the means of the two series and the standard deviations of their innovations can vary over time with the underlying regime, S_t. I constrain this regime to be the same across the two series and assume that this single latent variable (the state of the financial sector) plays a role in determining both observed series.

Following Hamilton, I model this regime as the outcome of an unobserved, discrete time, discrete state, first-order Markov process. There are two states: state zero is financial calm and state one is financial crisis. The means of the two time series and the standard deviations of their innovations are

$$\mu_i(S_t) = \alpha_{i,0} + \alpha_{i,1}S_t \text{ and } \sigma_i(S_t) = \omega_{i,0} + \omega_{i,1}S_t\,. \qquad (2)$$

For example, in a period of financial calm the mean of the yield spread is $\alpha_{2,0}$ while in a financial crisis it is $\alpha_{2,0} + \alpha_{2,1}$. The discussion earlier suggests that in the financial crisis state the growth rate of the deposit-currency ratio is negative and the yield spread is higher than it is in the state of financial calm. This implies that $\alpha_{1,0} + \alpha_{1,1} < 0$ and $\alpha_{2,1} > 0$. In each period there is a probability of switching between these two states. These time-invariant transition probabilities are

$$\begin{aligned} &P(S_t = 1|S_{t-1} = 1) = p\,; && P(S_t = 0|S_{t-1} = 1) = 1 - p\,; \\ &P(S_t = 1|S_{t-1} = 0) = 1 - q\,; && P(S_t = 0|S_{t-1} = 0) = q\,. \end{aligned} \qquad (3)$$

9. The ADF statistics for the test of the null hypothesis that a series is I(1) against the alternative that it is stationary around a constant are −3.45 and −3.05 for the growth rate of the deposit-currency ratio and the yield spread respectively. Both of these imply a rejection of the null hypothesis at the 5 percent level.

10. This is chosen using the Bayesian Information Criteria with a maximum lag length of 6. Using values of n ranging from 1 to 6 yields estimated time series of probabilities similar to those presented here and does not change any of the conclusions drawn. Equation (1) also imposes the restrictions that the cross-terms in the VAR matrices are zero and that the covariance matrix is diagonal. Again, relaxing these restrictions does not alter the time series of conditional probabilities reported in the following section.

Since the current regime and any past regimes are unobserved, inference about these regimes is based on the observable time series, in this case the growth rate of the deposit-currency ratio and the yield spread. This is done using the nonlinear filter described in Appendix 2. The filter outputs a time series of conditional probabilities over the event that current state is one of financial crisis, that is, $P(S_t = 1|\mathbf{x_t},\mathbf{x_{t-1}},\ldots,\mathbf{x_0})$. It also allows for the updating of previous conditional probabilities. In this case, with four lags, the conditional probability of financial crisis four periods previous can be updated to give probability, $P(S_{t-4} = 1|\mathbf{x_t},\mathbf{x_{t-1}},\ldots,\mathbf{x_0})$. Finally, the filter outputs a time series for the conditional likelihood that facilitates maximum likelihood estimation of the model's parameters.

To estimate the model described above I use the monthly data on the growth rate of the deposit-currency ratio and the yield spread from 1919 to 1941. The results from the maximum likelihood estimation of this model and a single-state model appear in Table 1. The Markov-switching model has six more parameters than the single-state model. A likelihood ratio test of the null hypothesis that the data were generated by the single-state model, against the alternative of the Markov-switching model, yields a test statistic of approximately 500, suggesting a comfortable rejection of the null hypothesis. However, under the null of a one-state model, the parameters describing the second state are unidentified. This means that the likelihood ratio test statistic does not possess the standard $\chi^2(6)$ distribution.[11]

TABLE 1

PARAMETER ESTIMATES FOR MARKOV-SWITCHING AND SINGLE-STATE MODELS

	Markov-Switching		Single-State	
	Estimate	Standard Error	Estimate	Standard Error
p	0.823	(0.072)	—	—
q	0.964	(0.014)	—	—
$\alpha_{1,0}$	0.204	(0.129)	−0.062	(0.219)
$\alpha_{1,1}$	−2.227	(0.767)	—	—
$\alpha_{2,0}$	2.484	(0.624)	2.891	(0.416)
$\alpha_{2,1}$	0.207	(0.050)	—	—
$\rho_{1,1}$	0.047	(0.104)	0.309	(0.061)
$\rho_{1,2}$	0.163	(0.101)	−0.020	(0.063)
$\rho_{1,3}$	0.200	(0.101)	0.145	(0.063)
$\rho_{1,4}$	−0.068	(0.101)	−0.065	(0.061)
$\rho_{2,1}$	1.184	(0.046)	1.195	(0.102)
$\rho_{2,2}$	−0.294	(0.611)	−0.211	(0.089)
$\rho_{2,3}$	0.027	(0.064)	−0.386	(0.089)
$\rho_{2,4}$	0.071	(0.145)	0.363	(0.057)
$\omega_{1,0}$	1.745	(0.088)	3.207	(0.137)
$\omega_{1,1}$	5.298	(0.873)	—	—
$\omega_{2,0}$	0.139	(0.012)	0.374	(0.016)
$\omega_{2,1}$	0.785	(0.113)	—	
Log-likelihood	−573.94		−821.40	

11. For the likelihood ratio test to have an asymptotic χ^2 distribution we require that the information matrix is nonsingular. If one tries to fit an Q-state model when the true process has $Q-1$-states this condition, does not hold. See, for example, Garcia (1998) or Coe (2000) for a more detailed discussion of this issue.

To tackle this problem I perform a Monte Carlo experiment to generate an empirical critical value for my sample test statistic. I create artificial data under the null hypothesis that the data-generating process is the single-state model described in Table 1.[12] I then use this artificial data to estimate both the single-state model and the Markov-switching model by maximum likelihood. Finally, I calculate the likelihood ratio test statistic for the test of the null hypothesis of the single-state model against the alternative of the Markov-switching model. I repeat this procedure one thousand times to obtain an empirical distribution for this test statistic. The percentiles of this distribution are similar to those of the asymptotic $\chi^2(6)$ distribution. For example, the empirical 95 percent and 99 percent critical values are 10.36 and 16.92, respectively. Clearly, these critical values are much lower than the sample test statistic, supporting the use of the Markov-switching model.

I also follow the suggestion of Hamilton (1988, 1994) and conduct a test on the one state model to see whether a two-state model is required. I use a Lagrange Multiplier test in which the null hypothesis of the constant parameter single-state model with homoskedastic errors is tested against an alternative in which the variance of the residuals depends on the lagged filter output. For both sets of residuals the null hypothesis that the lagged filter output contains no explanatory power can be comfortably rejected. Again this supports the use of the Markov-switching model. For the growth rate of the deposit-currency ratio the test statistic is 21.69. For the yield spread it is 38.71. Under the null hypothesis, both of these statistics are distributed $\chi^2(1)$.

Moving on to the parameter estimates, the estimated Markov probabilities give the unconditional probabilities of a change in regime. If the state at time $t-1$ was financial calm then there is a 96 percent chance that state at time t will also be financial calm. On the other hand, if the state at time $t-1$ was financial crisis, there is an 82 percent chance that the state at time t will also be financial crisis.

The estimated means of the growth rate of the deposit-currency ratio and the yield spread in the state of financial calm are 0.204 and 2.484, respectively. The former is insignificantly different from zero in the statistical sense. In the financial crisis state the estimated mean growth rate of the deposit-currency ratio falls to -2.025. This implies that during a financial crisis the deposit-currency ratio falls by an average of 2 percent per month. The estimated mean of the yield spread rises to 2.691. A test that the mean of a series is the same in the two regimes is just a simple test of the null hypothesis that the incremental parameter, $\alpha_{1,1}$ or $\alpha_{2,1}$ is equal to zero. For both series this test produces a rejection of the null hypothesis at the 1 percent level. The null hypothesis $\alpha_{0,1} + \alpha_{1,1} < 0$ is also rejected at the 1 percent level against the alternative $\alpha_{0,1} + \alpha_{1,1} > 0$. The negative growth rate of the deposit-currency ratio and higher yield spread are consistent with state one being a state of financial crisis.[13]

12. To do this I generate data using the parameter estimates for the linear model shown in Table 1 and two sets of random numbers. These random numbers are distributed $N(0,\omega_{i,0}^2)$. The standard deviations of the innovations to these series are $\omega_{1,0} = 3.207$ and $\omega_{2,0} = 0.374$, respectively. I generate series of length $T+200$ where $T=276$ is the sample size. I then discard the first 200 observations to minimize the influence of starting values.

13. The variance of the innovations to both series is higher in the financial crisis state and again these differences are statistically significant.

Moving on, I can use the parameter estimates from Table 1 and the two data series to calculate a conditional probability of financial crisis for each date in my sample. Figure 2a shows the conditional probability of financial crisis, $P(S_t = 1|\mathbf{x_t},\mathbf{x_{t-1}},\ldots,\mathbf{x_0})$ from May 1919 to December 1941. This probability is conditional on the two data series up to and including the current period and the parameter estimates in Table 1. As mentioned earlier, the filter allows for the updating of conditional probability over previous states. The updated conditional probabilities, $P(S_t = 1|\mathbf{x_{t+4}},\mathbf{x_{t+3}},\ldots,\mathbf{x_0})$ are shown in Figure 2b. This is the probability of financial crisis at time t, conditional on the two series up to and including period $t+4$ and the parameter estimates. Clearly, this series is much smoother than the one in Figure 2a.

If the disruptions to the financial sector during the Great Depression can be thought of as a shift to a regime of financial crisis, this should be reflected in the time series of conditional probabilities. More specifically, one would expect to see a probability of close to one assigned to the financial crisis regime during the early 1930s. The remainder of this section discusses the changes in regime implied by the conditional probabilities in Figure 2, in the context of the surrounding events.[14]

Figures 2a and 2b are consistent with Mishkin's (1991) assertion that a stock market crash alone does not necessarily imply a financial crisis. Following the crash of October 1929, the conditional probability of financial crisis jumps to one in November 1929. However, this lasts for only one month and by December 1929 is close to zero again. This evidence suggests that the financial crisis during the Great Depression does not begin with the stock market crash of 1929.

The probability of financial crisis then stays at approximately zero until October 1930. The estimated conditional probabilities suggest that a shift into the financial crisis state occurs then. This coincides with the first banking crisis and the time when Friedman and Schwartz argue that "a contagion of fear spread among depositors" (1963, p. 308). There is then a brief move back to the state of financial calm in January 1931. However, this is short-lived and, following the onset of the second banking crisis in March 1931, the conditional probability of financial crisis jumps back up in April 1931. From then on, with the exception of a short three-month period in late 1932, the updated conditional probability of financial crisis remains high until February 1934.[15] Here it is interesting to note that two attempts to alleviate the financial crisis during early 1932, the establishment of the RFC and the open market operations, have no effect on these conditional probabilities.

Perhaps the most interesting feature of Figures 2a and 2b is the implication they have for the ending of the financial crisis. If the reforms discussed in the previous section did have a positive effect on the financial system, this should be reflected in the time series of estimated conditional probabilities over the current state of the financial system. Given that the early 1930s is a period of financial crisis, one would expect the crisis to end in 1933 or 1934. Some combination of the traditional view

14. The estimated conditional probabilities suggest a short crisis in October 1921. This only lasts for one month and is not discussed here. The evidence for this month of financial crisis is greatly weakened when other indicators of financial crisis are considered.

15. This period is September to November 1932.

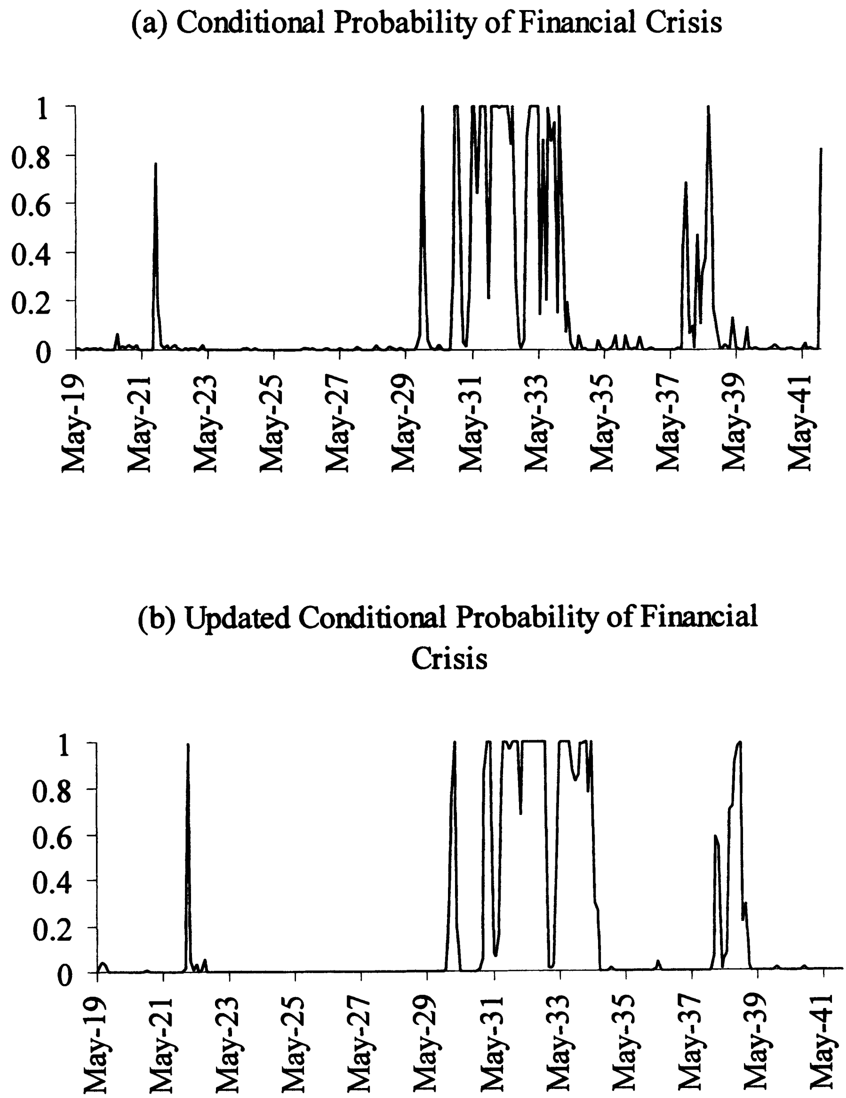

FIG. 2. Conditional Probability of Financial Crisis, 1919 M5–1941 M12

and Wigmore's view suggests that there would be a regime change in the spring of 1933. On the other hand, the view that the introduction of the FDIC ended the financial crisis dates the regime change as being in early 1934.

Figure 2a shows a fall in the conditional probability of financial crisis to 0.15 in May 1933. However, this is temporary. For the majority of 1933 the probability of financial crisis remains above 0.8, suggesting no change in regime immediately following the reforms of the Spring of 1933. This point is emphasized by looking at the updated probabilities in Figure 2b. This updated probability of financial crisis for May 1933 is 0.88. In fact, for the whole of 1933 it is never below 0.78. This suggests that while the reforms contained in the EBA and the abandonment of the gold standard may have been necessary, they were not sufficient to end to the financial crisis.

The conditional probabilities suggest that the financial crisis ends in February of 1934. This is immediately after the introduction of the FDIC in the previous month and the sharp increase in authorized lending by the RFC in December 1933 and January 1934. This is shown more clearly in the updated probabilities of Figure 2b than those in Figure 2a. Figure 2b shows a probability of financial crisis of 0.301 in February 1934. This falls to 0.266 in March and is zero for the remainder of 1934. This result is consistent with the view that at least one of the introduction of the FDIC and the increased lending by the RFC was crucial for ending the financial crisis.

During the first quarter of 1934 the introduction of the FDIC was not the only major policy change. On the 31st January 1934, Roosevelt announced that the dollar price of gold was to be fixed at $35 per fine ounce. While this did not represent a devaluation of the dollar (it had already depreciated to this level by July 1933), it did represent the potential for a significant expansion of high-powered money as the Treasury revalued its gold holdings. Previously these holdings were valued at $20.67 per fine ounce; now the Treasury valued them at $35 per fine ounce. As a result, it could print additional paper money in the form of gold certificates that had a nominal value of nearly $3bn. This could be done without the acquisition of additional gold to maintain backing. This might lead one to argue that monetary expansion ended the financial crisis.[16] However, data on high-powered money show that the potential for monetary expansion was not realized in early 1934. The increase that did occur was only sufficient to restore the stock of high-powered money to approximately its February 1933 level.[17]

Figures 2a and 2b also suggest a brief period of financial crisis in late 1937 and early 1938. Between 1936 and 1937 the Federal Reserve more than doubled the required reserve ratio for member banks from 6.2 percent of total assets to 12.6 percent of total assets.[18] An increase in required reserves can lead to a fall in the amount of loans that banks will make. Therefore, as the banking system reduces its role as a credit interme-

16. Friedman and Schwartz argue that during the Great Depression the Federal Reserve should have, but did not, provide a significant expansion of high-powered money to protect the banking sector.

17. In February 1933 the stock of high-powered money was $8.807bn, in March 1934 it was $8.998bn (Friedman and Schwartz 1963, Appendix B). These are nominal figures. The growth in the real stock of high-powered money was even lower due to rising prices after March 1933.

18. Friedman and Schwartz (1963, Table 19).

diary the cost of credit intermediation rises. As a result, it is possible that by raising reserve requirements the Federal Reserve would have caused the cost of credit intermediation to move as it would during a financial crisis. In order for this to happen it must be the case that banks wished to keep their level of excess reserves constant following the change in required reserves, as Friedman and Schwartz suggest:

> When the rise in reserve requirements immobilized the accumulated cash, they (member banks) proceeded rather promptly to accumulate additional cash for liquidity purposes. (Friedman and Schwartz 1963, p. 458)

3. FINANCIAL CRISIS AND OUTPUT GROWTH

In this section, I use lagged estimated conditional probabilities of financial crisis as explanatory variables for output growth during the interwar period. I show that including these probabilities in a reduced-form equation provides additional explanatory power for output fluctuations during that period. This is explanatory power in addition to that provided by lagged values of output, the money stock, the growth rate of the deposit-currency ratio and the yield spread.

I estimate the following equation using monthly data from August 1919 to December 1941 by OLS:

$$\Delta y_t = \beta_0 + \sum_{k=1}^{n} \beta_k \Delta y_{t-k} + \sum_{k=1}^{n} \gamma_k \Delta m_{t-k} + \sum_{k=1}^{n} \delta_{1,k} x_{1,t-k} + \sum_{k=1}^{n} \delta_{2,k} x_{2,t-k}$$
$$+ \sum_{l=1}^{5} \lambda_l P(S_{t-l} = 1 | \mathbf{x}_{t-1}, \mathbf{x}_{t-2}, \ldots, \mathbf{x}_0) + e_t \,. \tag{4}$$

Here Δy_t is the one-month change in the natural logarithm of manufacturing production, Δm_t is the one-month change in the natural logarithm of the money stock, $x_{1,t}$, $x_{2,t}$, and $\mathbf{x}_t$ are as described earlier and $e_t \sim i.i.d.N(0,\sigma_e^2)$. Under the null hypothesis that the state of the financial sector contains no additional explanatory power for output growth, $\lambda_l = 0$ for all $l = 1,2,\ldots,5$. The test statistic for this hypothesis is distributed $F(5,269 - K)$ under the null hypothesis, where K is the total number of estimated parameters in the unrestricted model.

I estimate this equation three times. In the first case the null model is restricted to contain only a constant and lagged values of the growth rate of manufacturing production. In other words, $\gamma_k = \delta_{1,k} = \delta_{2,k} = 0$ for all $k = 1,2,\ldots,n$. The question of interest is whether the financial crisis provides explanatory power for output fluctuations, independent of output's own history. In the second case lagged values of the money stock are added to the equation. That is, the $\gamma_k = 0$ for all $k = 1,2,\ldots,n$ restriction is relaxed. Here I am interested in Bernanke's proposition that the financial crisis had nonmonetary as well as monetary effects. Finally, the model is estimated without any maintained restrictions. Here I test the null hypothesis that the regime shifting effects contain no additional explanatory power above the linear measures of financial crisis already present.

The results from the estimation of this equation are reported in Table 2 and are for $n = 2$.[19] The first row shows the case where the null model contains only past values of the dependent variable. Here, the null hypothesis is that the state of the financial sector has no explanatory power for output fluctuations, beyond output's own history. The $F(5,261)$ statistic of 3.44 implies a rejection of this null hypothesis at the 1 percent level. This is consistent with the hypothesis that the state of the financial sector contains information useful for explaining output fluctuations. To obtain an approximate measure of the importance of the financial crisis, I use this estimated equation to calculate an average growth rate of output associated with each of the two states. This suggests that periods of financial crisis are associated with an annual growth rate of output of approximately −5.7 percent, compared to approximately 4.5 percent in periods of financial calm.[20]

As discussed above, Bernanke proposes that the financial crisis had additional effects beyond those that occurred through a monetary channel. The second row of Table 2 provides empirical support for this proposition. It shows that when lagged values of the growth rate of M2 are added to the equation for manufacturing production growth the null hypothesis is still rejected at the 1 percent level. In this case the $F(5,259)$ statistic is 3.19. This is consistent with the view that the financial crisis had nonmonetary effects along the lines suggested by Bernanke as well as the monetary effects emphasized by Friedman and Schwartz.

TABLE 2

F-TESTS OF THE NULL HYPOTHESIS $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5 = 0$ IN EQUATION (4)

Maintained Restrictions	*F*-Statistic	$T - K$	*p*-value	$\sum_{l=1}^{5}\lambda_l$	(s.e.)
$\gamma_1 = \gamma_2 = 0$ $\delta_{1,1} = \delta_{1,2} = 0$ $\delta_{2,1} = \delta_{2,2} = 0$	3.436	261	0.005	−0.603	(0.857)
$\delta_{1,1} = \delta_{1,2} = 0$ $\delta_{2,1} = \delta_{2,2} = 0$	3.187	259	0.008	−0.584	(0.985)
None	3.002	255	0.012	−1.510	(1.408)

19. For all values of n from 1 to 6 the null hypothesis can be rejected at least at the 5 percent level in each of the three experiments. $n = 2$ is the number selected by sequentially testing down using an AR model for manufacturing production growth.

20. These numbers should be treated with caution, as the confidence intervals around these sums are quite wide as the standard errors for $\sum_{l=1}^{5}\lambda_l$ in Table 2 suggest. At first glance this might appear to be inconsistent with the F-statistics in Table 2. However, not all $\hat{\lambda}_l$ are negative. The estimates of p and q imply that as l increases, the unconditional probability that period t is one of crisis given $t-l$ was one of crisis falls to a number in the region of 0.5 quite quickly. Therefore, it is unsurprising that among the higher values of l there is a positive estimate of λ_l.

Finally, the work of Cooper and Ejarque (1995) and Cooper and Corbae (1999) suggests modeling the financial sector as moving between an optimistic and a pessimistic steady state during the Great Depression. The evidence pointing to the presence of regime shifts in the previous section is consistent with this view. Here I present evidence to suggest that these regime shifts have additional explanatory power for output fluctuations above linear measures of the crisis. When lagged values of the growth rate of the deposit-currency ratio and the yield spread are added to the equation, the null hypothesis that the conditional probabilities contain no additional explanatory power is still rejected at the 5 percent level. In this case the $F(5,255)$ statistic is 3.00. This suggests that modeling the financial crisis as a regime shift can help explain output fluctuations in the interwar period.

4. CONCLUSIONS

In this paper I explore the timing and effects of the 1930s' financial crisis in the United States using a regime-switching approach. Temin and Wigmore (1990) argue that the New Deal represented a change in policy regime that played a crucial role in ending the Great Depression. I find evidence consistent with such a regime change occurring in the financial sector. Estimated probabilities over the state of the financial sector suggest that the financial crisis did not begin with the 1929 stock market crash, but with the first banking panic in late 1930.

Contrary to the traditional view that the financial crisis ended with the bank holiday in March 1933, or Wigmore's view that it ended with the decision to leave the gold standard, my results suggest that a regime change signaling the end of the financial crisis did not occur until early 1934. This result is consistent with the view that some combination of the introduction of the FDIC in January 1934 and increased RFC loan activity was crucial to restoring stability to the financial sector. The implication for modern day crises, such as that in southeast Asia, is that the more successful policies for ending a crisis may be those directed at reforming domestic institutions. In particular a willingness of the government to protect depositor wealth appears to be important.

I also present evidence to show that the state of the financial sector contains significant explanatory power for output fluctuations during the interwar period. These results suggest that the expected growth of manufacturing industrial production was lower during a financial crisis. I also find evidence to suggest that the effects of the financial crisis are not confined to the monetary channel. This bolsters Bernanke's view that the financial crisis also contributed to the contraction through a nonmonetary channel. Finally, I find that the conditional probability of financial crisis retains its explanatory power for output fluctuations when linear measures of the financial crisis are added to the empirical model. This result suggests that the incorporation of regime shifts in the financial sector may help improve our understanding of the Great Depression and other financial crises.

APPENDIX ONE: DATA SOURCES

Deposit-Currency Ratio:	Friedman and Schwartz (1963, Table B-3)
Commercial Bank Deposits ($m.):	Friedman and Schwartz (1963, Table A-1) Federal Reserve Bulletin (various issues)
Yield on Baa-Rated Bonds (percent p.a)	Banking and Monetary Statistics (1943,Table 129)
Yield on Aaa-Rated Bonds (percent p.a.):	Banking and Monetary Statistics (1943,Table 129)
Yield on Long-term Government Bonds (percent p.a.):	Banking and Monetary Statistics (1943, Table 129)
Consumer Price Index (1957–59 = 100):	NBER Macrohistory Database (series M04128)
Industrial Production (Manufacturing):	NBER Macrohistory Database (series M01175)
Money Stock ($bn):	Friedman and Schwartz (1963, Table A-1)

Real bank deposits and the real liabilities of suspended banks are calculated by deflating the appropriate nominal series by the consumer price index. Monthly growth rates are calculated as $100^*\ln(z_t/z_{t-1})$ for z equal to the deposit-currency ratio, bank deposits, real bank deposits, manufacturing production, and the money stock.

A potential problem with the data on liabilities in suspended banks is how to treat the bank holiday announced on March 6th. The observation for March 1933 for this series is seven times as large as the next highest observation. I deal with this observation in the same way as Bernanke (1983) and multiply it by 0.15. This means that the suspensions that occurred under government control in March 1933 are treated as similar in magnitude to those that occurred in October 1931 without intervention.

APPENDIX TWO: FILTER FOR MARKOV-SWITCHING MODEL

This appendix describes the filter used in the estimation of the Markov-switching model. Here, $P(z)$ denotes $P(Z = z)$ when z is discrete and density function $f(z)$ when z is continuous.

Step One: Using the Markov transition probabilities calculate

$$P(s_t,s_{t-1},\ldots,s_{t-n}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) = P(s_t|s_{t-1}) \times P(s_{t-1},\ldots,s_{t-n}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) \,,$$

where the second term on the right-hand side is the output from the previous iteration of the filter.

Step Two: The joint conditional density of $\mathbf{x_t}$ and $(S_t,S_{t-1},\ldots s_{t-n})$ is

$$P(\mathbf{x_t},s_t,s_{t-1},\ldots,s_{t-n}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) = P(\mathbf{x_t},|s_t,s_{t-1},\ldots,s_{t-n},\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) \\ \times P(s_t,s_{t-1},\ldots,s_{t-n}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) ,$$

where

$$P(\mathbf{x_t}|s_t,s_{t-1},\ldots,s_{t-n},\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) = \frac{1}{2\pi\sigma_1(S_t)\sigma_2(S_t)} \\ \times \exp\left\{\sum_{i=1}^{2} -\frac{1}{\sigma_i^2(S_t)}\begin{bmatrix} x_{i,t} - \mu_i(S_t) - \rho_{i,1}\left(x_{i,t-1} - \mu_i(S_{t-1})\right) \\ -\ldots-\rho_{i,n}\left(x_{i,t-n} - \mu_i(S_{t-n})\right) \end{bmatrix}^2\right\}.$$

Step Three: The branch of the likelihood function for observation t is

$$P(\mathbf{x_t}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) = \sum_{s_t=0}^{1}\cdots\sum_{s_{t-n}=0}^{1} P(\mathbf{x_t},s_t,s_{t-1},\ldots,s_{t-n}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0}).$$

Step Four: Steps two and three combined give

$$P(s_t,s_{t-1},\ldots,s_{t-n}|\mathbf{x_t},\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) = \frac{P(\mathbf{x_t},s_t,s_{t-1},\ldots,s_{t-n}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0})}{P(\mathbf{x_t}|\mathbf{x_{t-1}},\ldots,\mathbf{x_0})}.$$

Step Five: The output of iteration t is

$$P(s_t,s_{t-1},\ldots,s_{t-n+1}|\mathbf{x_t},\mathbf{x_{t-1}},\ldots,\mathbf{x_0}) = \sum_{s_{t-n}=0}^{1} P(s_t,s_{t-1},\ldots,s_{t-n}|\mathbf{x_t},\mathbf{x_{t-1}},\ldots,\mathbf{x_0}).$$

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