Taylor's Rule and the Fed: 1970–1997

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This paper estimates a simple model of the Federal Reserve's "reaction function"—that is, the relationship between economic developments and the Fed's response to them. We focus on how this estimated reaction function has changed over time. Such changes are not surprising given compositional changes in the Federal Open Market Committee, and we consider three subsamples delineated by the terms of recent Fed Chairmen. We find that the estimated reaction functions for each period vary in ways that seem broadly consistent with the success or failure during the period at controlling inflation. These results suggest that a Taylor-rule framework is a useful way to summarize key elements of monetary policy.

Macroeconomists have long been interested in modeling the Federal Reserve's "reaction function"—that is, modeling how the Fed alters monetary policy in response to economic developments. The Fed's reaction function plays an important role in a wide variety of macroeconomic analyses. It can provide a basis for forecasting changes in the Fed's policy instrument—namely, short-term interest rates. Also, within the context of a macro model, the reaction function is an important element in evaluating Fed policy and the effects of other policy actions (e.g., fiscal policy) or economic shocks (e.g., the 1970s oil embargo). Finally, when rational expectations are assumed in macro models, knowing the correct reaction function is an important element in estimating the entire model. For example, with forward-looking expectations, estimates of a parameter such as the one linking real spending to the policy instrument will likely depend crucially on expected monetary policy and the nature of the monetary policy regime.

Numerous reaction functions have been estimated. Khoury (1990), for example, surveys 42 such empirical Fed reaction functions from various studies. Moreover, the large numbers of monetary policy vector autoregressions (VARs) (e.g., Bernanke and Blinder 1992) that have been estimated recently also include an equation that can be interpreted as a Fed reaction function (see Rudebusch 1998).

However, despite this work, researchers have not been particularly successful in providing a definitive representation of Fed behavior. Khoury finds little consistency in the significance of various regressors in the reaction functions she surveys. She states (p. 28): "One who [examines] just one of these reaction functions may feel convinced that one has learned how the Fed responds to economic conditions, but that seeming knowledge disappears as one reads a large number of these studies." Overall, it appears that there have not been any great successes in modeling Fed behavior with a single, stable reaction function. As an illustration, McNees (1992) compares his latest estimate of a Fed reaction function to his previous estimate (1986) and states (p. 11): "The number of modifications to the original specification required to make it track the past six years serve[s] as a clear illustration that policy reaction functions can be fragile."

There are a number of plausible explanations for such instability. For example, a central bank's reactions may be too complex to be adequately captured by a simple linear regression. Another factor may be changes in the composition of the Federal Open Market Committee (FOMC). Such compositional changes may bring to the fore policymakers with different preferences and different conceptions of the appropriate operation and likely transmission of monetary policy. While many people and events influence policy, arguably one of the more important and *identifiable* compositional changes is in the Fed Chairmanship. Changes associated with different Chairmen may be exogenous, but there also may be an endogenous element that represents an adaption to "lessons" learned from prior experiences. Indeed, Chairmen may be chosen who are seen as likely to avoid the mistakes of the past.[1] For example, part of the backing for Paul Volcker as Chairman in the high-inflation environment of 1979 may have come from the expectation that he would be tough on inflation.[2]

Accordingly, in this paper we estimate reaction functions for three separate empirical subsamples delineated by the identity of the Fed Chairman: specifically, we consider the terms of Arthur Burns (1970.Q1–1978.Q1), Paul Volcker (1979.Q3–1987.Q2), and Alan Greenspan (1987.Q3 –present). We omit any discussion of policy under Chairman Miller (1978.Q2–1979.Q2) because of his short tenure. This delineation gives us three subsamples of moderate and approximately equal length that are selected in an *a priori* fashion.

The organizing principle for our investigation is the Taylor rule, which we use as a rough gauge for characterizing and evaluating the broad differences in the relative weights given to monetary policy goal variables between periods. The rule specifies that the real federal funds rate reacts to two key goal variables—deviations of contemporaneous inflation from an inflation target and deviations of real output from its long-run potential level. These variables would appear to be consistent with the Fed's legislated mandate.[3] Moreover, Taylor (1993) argued that this rule represents "good" policy, in the sense that it relates a plausible Fed instrument to reasonable goal variables, and it stabilizes both inflation and output reasonably well in a variety of macroeconomic models. More recent model simulation studies (e.g., Rudebusch and Svensson 1998 and Levin, Wieland, and Williams 1997) have reinforced the latter conclusion.

Moreover, these recent studies suggest that although Taylor-type rules are very simple, they may be capable of capturing the essential elements of more realistic regimes in which the central bank "looks at everything." Simple Taylor-type reaction functions were found to perform almost as well as optimal, forecast-based reaction functions that incorporate all the information available in the models examined. In addition, the simple specification was found to perform almost as well as reaction functions that explicitly include a variety of additional variables. These results appear to be fairly robust across a variety of macroeconomic models. Thus, the general form of the Taylor rule may be a good device for capturing the key elements of policy in a variety of policy regimes.

The rule as originally specified by Taylor serves as a useful starting point for our investigation below. After briefly examining this rule, we focus on econometrically estimating a dynamic version of the rule for the three periods defined above. We find that, overall, the estimated dynamic Taylor-type reaction functions do provide a way to capture important elements of the policy regimes in place during these periods. The key elements of the estimated reaction functions for each period also vary in ways that seem broadly consistent with the success or failure during the periods at controlling inflation. This conclusion is reinforced at the end of the paper by explicit evaluations of the reaction functions in the three periods in the context of a small macro model.

We do not regard the results of this investigation as providing a complete representation of Fed behavior, in part because we have controlled for only one source of sample instability. However, we hope our results serve as a springboard for a discussion of some of the salient features—and changes—in Federal Reserve behavior over time. Also, as noted above, it is important to develop a better understanding of how the Fed's reaction function has changed over time for macroeconometric research.

1. However, attempts to avoid the mistakes of the past sometimes may lead to new mistakes. De Long (1997, p. 250) argues that ". . . at the deepest level, the truest cause of the inflation in the 1970s was the shadow cast by the great depression . . . ."

2. De Long (1997, p. 274) argues, "A mandate to fight inflation by inducing a significant recession was in place by 1979, as a result of a combination of fears about the cost of inflation, worry about what the 'transformation of every business venture into a speculation on monetary policy' was doing to the underlying prosperity of the American economy, and fear that the structure of expectations was about to become unanchored and that permanent double-digit inflation was about to become a possibility."

3. The 1977 amendment to the Federal Reserve Act requires the Fed to "promote effectively the goals of maximum employment, stable prices, and moderate long-term interest rates." The Humphrey-Hawkins Act of 1978 affirms the responsibility of the federal government in general to promote "full employment and production, . . . and reasonable price stability," among other things.

I. TAYLOR'S ORIGINAL RULE

Taylor (1993) suggests a very specific and simple rule for monetary policy. His rule sets the level of the nominal federal funds rate equal to the rate of inflation (in effect, making it an equation for the ex post *real* funds rate) plus an "equilibrium" real funds rate (a "natural" rate that is seen as consistent with full employment) plus an equally weighted average of two gaps: (1) the four-quarter moving average of actual inflation in the GDP deflator less a target rate, and (2) the percent deviation of real GDP from an estimate of its potential level:

(1) $$i_t = \pi_t + r^* + 0.5(\pi_t - \pi^*) + 0.5(y_t)$$

where i = federal funds rate,
r^* = equilibrium real federal funds rate,
π = average inflation rate over the contemporaneous and prior three quarters (GDP deflator),
π^* = target inflation rate
y = output gap (100·(real GDP – potential GDP) ÷ potential GDP).[4]

Taylor did not econometrically estimate this equation. He assumed that the weights the Fed gave to deviations of inflation and output were both equal to 0.5; thus, for example, if inflation were 1 percentage point above its target, the Fed would set the real funds rate 50 basis points above its equilibrium value. Furthermore, Taylor assumed that the equilibrium real interest rate and the inflation target were both equal to 2 percent. We shall examine these assumptions below; however, it is instructive to consider the interest rate recommendations from the original Taylor rule.[5]

Figure 1 illustrates the original Taylor rule during 1970–1998. The top panel shows the recommendations of the rule on a quarterly basis. The bottom two panels show the variables that enter the rule—the GDP gap and inflation. As explained earlier, higher levels of both variables lead to a higher level of the recommended funds rate. In 1979, for example, the rule recommended a high funds rate mainly because inflation was quite high, and to a lesser extent, because real GDP exceeded its potential level by a small amount.

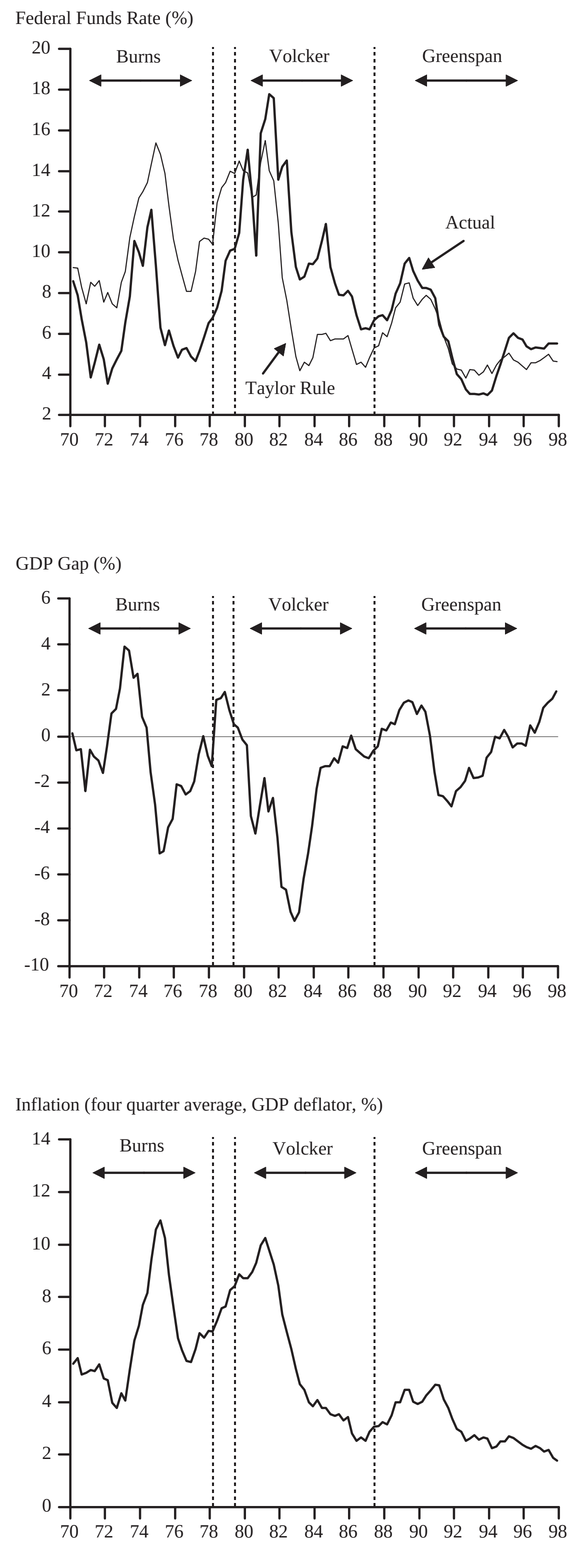
As shown in Figure 1, the original Taylor rule fits reasonably well to the actual funds rate during the Greenspan period. It captures the major swings in the funds rate over the period, but with less amplitude. The R^2 for the period is 87 percent for quarterly levels of the nominal funds rate,

4. Taylor (1993) used a log linear trend of real GDP over 1984.Q1 to 1992.Q3 as a measure of potential GDP. As discussed below, we have used a more flexible structural estimate.

5. For a complementary analysis, see Taylor (1997).

FIGURE 1

THE TAYLOR RULE AND ITS COMPONENTS



and 52 percent for quarterly changes. The arguments in the rule—inflation and the GDP gap—roughly correspond with goals legislated for U.S. monetary policy—stable prices and full employment. In this spirit, Governor Meyer (1998) stresses that stabilizing real GDP around its trend in the short run and controlling inflation in the longer term are important concerns of the Fed. Although U.S. policymakers look at many economic and financial indicators, the two gaps specified in the rule may be highly stylized measures of important short- and long-run concerns. Also, the GDP gap can be interpreted not only as a measure of business cycle conditions, but also as an indicator of future inflation in the context of a Phillips curve model. Measures of the productive capacity of the U.S. economy, whether measured by potential GDP, industrial capacity, or the "natural" rate of unemployment, appear to figure prominently in Fed forecasts of future inflation (Greenspan 1995).[6] Overall, by focusing on policy responses to the Fed's basic goal variables, the Taylor rule implicitly captures policy responses to the many economic factors that affect the evolution of those goal variables.[7]

Judd and Trehan (1995) argue that the Taylor rule also provides some perspective on policies during the Burns and Volcker periods. With regard to the Burns period, although the movement of the actual funds rate was highly correlated with the rule's prescriptions, the funds rate itself was consistently lower than the rule's recommended rate (Figure 1). This result is consistent with the overall increase in inflation during this period, and it confirms that the rule, with its explicit 2 percent inflation target, might have held inflation to a much lower level than policy actually did. During the Volcker period, when the Fed significantly reduced inflation, the funds rate was consistently higher than what the rule recommended, suggesting that the Fed was more aggressive in reducing inflation than the rule would have been.

II. Issues in Estimating Taylor's Rule

The original Taylor rule appears to provide a rough description of policy during the Greenspan period, as well as a useful benchmark for discussing the policy regimes in place during the Burns and Volcker periods. While the original rule provides a reasonable starting point, this section examines alternatives to Taylor's simple specification by econometrically estimating the reaction function weights, rather than simply choosing parameters as Taylor did. Estimating Taylor-type equations may provide a better description of Fed policy. We consider several issues in estimating a reaction function based on the Taylor rule, including the specification of dynamics, the equilibrium real rate, the inflation target, and the output gap.[8]

Estimating the Taylor Rule with Dynamics

It is fairly straightforward to estimate a Taylor rule as in equation (1). Simply replace the rule-based recommended nominal funds rate with the historical series, add a residual error term to capture deviations from the rule, and estimate the weights as coefficients. One complication to this procedure is that central banks often appear to adjust interest rates in a gradual fashion—taking small, distinct steps toward a desired setting (see, e.g., Rudebusch 1995). We allow for such interest rate smoothing by estimating the Taylor rule in the context of an error correction model. This approach allows for the possibility that the funds rate adjusts *gradually* to achieve the rate recommended by the rule.[9]

In our specification, we replace equation (1) with

$$(2) \qquad i_t^* = \pi_t + r^* + \lambda_1 (\pi_t - \pi^*) + \lambda_2 y_t + \lambda_3 y_{t-1}$$

where i_t^* is explicitly denoted as the recommended rate that will be achieved through gradual adjustment. Also, equation (2) includes an additional lagged gap term along with the contemporaneous gap. This is a general specification that allows for the possibility that the Fed responds to a variety of variables proposed as reasonable monetary policy targets, including inflation alone ($\lambda_2 = \lambda_3 = 0$, as in

6. Given the lags in the monetary transmission mechanism, an explicitly forward-looking version of the Taylor rule—with inflation and output *forecasts* as arguments—also might be appropriate. Clarida, Gali, and Gertler (1997a, 1997b) estimate a rule using inflation forecasts and obtain results similar to our own, and Rudebusch and Svensson (1998) examine the theoretical properties of such a rule.

7. The Taylor rule has gained the attention of some Fed policymakers (Blinder 1996, *Business Week* 1996, Meyer 1998, and Yellen 1996), who have used it as a helpful, broad characterization of U.S. monetary policy. In addition, it has gained some acceptance outside the Fed as a way to think about how the Fed might react to economic and inflationary developments (Prudential Economics 1996, and Salomon Brothers 1995a, 1995b). Of course, there are always questions about the reliability of any current implications of the rule because of uncertainty about the level of potential GDP. Some analysts argue that increased productivity, due to computer and other technological developments, means that potential output is being mis-measured. See Trehan (1997) for a discussion of the debate about productivity.

8. We use current data throughout this paper. It would be preferable to use the original data that policymakers actually were looking at when decisions about the funds rate were being made. Unfortunately, we do not have access to these data for our full 1970–1997 sample period. See Orphanides (1997) for an analysis of the effects of original versus final data in estimating a Taylor rule for the 1987–1992 period.

9. Mehra (1994) employs a similar dynamic specification.

Meltzer 1987), nominal GDP growth ($\lambda_1 = \lambda_2 = -\lambda_3$, as in McCallum 1988), inflation and real GDP growth with different weights ($\lambda_1 \neq \lambda_2 = -\lambda_3$), as well as inflation and the GDP gap in level form (as in Taylor 1993).

The dynamics of adjustment of the actual level of the funds rate to i_t^* are given by

$$\Delta i_t = \gamma(i_t^* - i_{t-1}) + \rho\Delta i_{t-1} \,. \tag{3}$$

That is, the change in the funds rate at time t partially corrects the "error" between last period's setting and the current recommended level (the first term), as well as maintaining some of the "momentum" from last period's funds rate change (the second term).[10]

By substituting equation (2) into (3), we obtain the equation to be estimated:

$$\Delta i_t = \gamma\alpha - \gamma i_{t-1} + \gamma(1 + \lambda_1)\pi_t + \gamma\lambda_2 y_t + \gamma\lambda_3 y_{t-1} + \rho\Delta i_{t-1} \,, \tag{4}$$

where $\alpha = r^* - \lambda_1\pi^*$. This equation provides estimates of the weights on inflation and output in the rule and on the speed of adjustment to the rule.

Determining r^* and π^*

As is clear from equation (4), our estimation cannot pin down *both* the equilibrium real funds rate (r^*) and the inflation target (π^*) simultaneously. These two terms are combined in the constant term (α) and cannot be identified separately. The economics of this lack of identification are clear in the original Taylor rule of equation (1): The contemporaneous arithmetic effect on the recommended policy rate is the same for a 1 percentage point increase in r^* and for a 2 percentage point decrease in π^*. If both of these magnitudes are unknown, then neither can be individually identified from the estimate of the single parameter α.

Of course, if we assume a particular value for the equilibrium funds rate, then, through the estimates of α and λ_1, we can obtain an estimate of the inflation target. Conversely, an assumption about the inflation target can yield an estimate of the equilibrium rate. Table 1 sheds some light on plausible estimates of these quantities. One simple benchmark for the equilibrium real funds rate is the average real rate that prevailed historically over periods with a common start and end inflation rate.[11] As shown in the first column of Table 1, over a long sample from the early 1960s to the present, inflation edged up, on net, only slightly, while the real funds rate averaged 2.39 percent, which appears to be in the range of reasonable estimates.[12] During the Greenspan period (column 2), the real rate averaged 2.82, which is a bit higher. However, this higher level is consistent with the fact that inflation fell more than 1 percentage point during the Greenspan sample while it rose slightly during the long sample. Certainly, Taylor's (1993) suggestion that 2 percent was a reasonable guess for the value of the equilibrium rate during the Greenspan period seems plausible. It is more difficult to pin down the equilibrium real rate in the Volcker period. During this period, the real

10. We think that this "error correction" framework is a useful one for the consideration of dynamics. However, although the funds rate, the output gap, and the inflation rate are highly persistent, we make no claims that they are nonstationary (consistent with Rudebusch 1993).

11. This is analogous to using the average unemployment rate over periods with no net change in inflation to estimate a constant "natural" rate of unemployment (or NAIRU).

12. The real rates in Table 1 are calculated on an ex post basis as in equation (1), but similar results were obtained using ex ante rates constructed with the one-year-ahead inflation forecasts from the Philadelphia Fed's inflation expectations survey.

TABLE 1

INTEREST RATES AND INFLATION

	LONG SAMPLE (61.Q1–97.Q4)	GREENSPAN (87.Q3–97.Q4)	VOLCKER (79.Q3–87.Q2)	BURNS (70.Q1–78.Q1)
Average real interest rate (%)	2.39	2.82	5.35	0.02
Percentage point change in inflation	0.38	–1.32	–5.81	1.23
Average inflation (%)	4.38	3.03	5.35	6.47
End-of-sample inflation (%)	1.77	1.77	3.07	6.69

NOTE: The change in inflation (in percentage points) is calculated as the difference in four-quarter inflation from the first quarter to the last quarter of the sample. End-of-sample inflation is average inflation over the final four quarters of the sample. Inflation is measured as the four-quarter change in the GDP deflator, and the interest rate is the federal funds rate.

rate averaged over 5 percent, but there was also a large decline in inflation, so this average rate is likely much higher than the equilibrium real rate.[13] Conversely, real rates averaged about zero during the Burns period, but during this time inflation and inflationary pressures were rising, so the equilibrium rate was most likely higher than the average.

It is less clear how to obtain implicit inflation targets from the historical data. Table 1 provides the average levels of inflation over the various samples. However, given the persistence of inflation, the assumption that the target inflation rate of policymakers has been achieved on average in various samples seems less plausible than the assumption regarding the real funds rate (i.e., that cyclical fluctuations have averaged out over time). Policymakers, notably in the early part of the Volcker period, could have "inherited" persistent inflation rates much different from their own target rate, which could then skew their sample averages. More interesting perhaps is the end-of-sample inflation rate, which gives a reading on what policymakers were able to achieve by the end of their tenure. Note that this rate for Greenspan is close to the 2 percent target assumed in Taylor (1993).

As this discussion should make clear, there is much uncertainty in choosing values for r^* and π^*.[14] Therefore, we will show results below under a variety of assumptions about these magnitudes.

Estimating Potential Real GDP

One final issue to consider is the specification of the real output gap, which is defined as the percentage difference between real GDP and potential GDP. Potential output is unobserved and must be estimated. An atheoretical method to do this is to fit a trend to the data—again, on the assumption that over time cyclical fluctuations average out. For example, Taylor (1993) simply used a linear trend of log real GDP over a short sample period (1984–1992) as a proxy for potential output. One also could use a segmented linear trend (following Perron 1989) or a quadratic trend (as in Clarida, Gertler, Gali 1997a, 1997b) or other nonstructural methods (see Cogley 1997).

We believe a structural approach to estimating the output gap is more appropriate conceptually than an atheoretic approach, since the presence of output in the policy rule not only may reflect an interest in stabilizing real fluctuations but also may provide policymakers some information on future inflation. The structural approach is also the one typically used by policymakers at the Fed and elsewhere. In this paper, we use a structural definition of potential GDP that was developed at the Congressional Budget Office (1995). It is denoted Y^*, and the associated gap is shown in Figure 1.[15] This measure of potential output is not a simple fitted GDP trend, but is estimated in terms of a relationship with future inflation similar to the way a time-varying NAIRU is estimated within the context of a Phillips curve. We examine the robustness of our results to alternative measures of the gap in the Appendix.

III. ESTIMATES OF REACTION FUNCTIONS

Our main hypothesis is that taking account of changes in Fed Chairmen helps to account for changes in the Fed's reaction function. Accordingly, we conduct Chow tests on equation (4) for two breaks during the 1970.Q1–1997.Q4 period corresponding to the terms of Chairmen Burns, Volcker, and Greenspan (the Miller term, 1978.Q2–1979.Q2, was excluded). While a finding of significant breaks in the data would not be strong evidence in favor of our hypothesis, it would at least be a reasonable initial step that should be taken before proceeding to estimate separate reaction functions for those periods.

The test gives the null hypothesis of no structural change a p-value of 0.00 (i.e., it rejects stability at the 0.00 percent level of significance). In addition, we looked at the Burns/Volcker period and tested for a break between their terms, and similarly at the Volcker/Greenspan period and tested for a break between their terms. These tests rejected stability at significance levels of 0.00 and 0.07, respectively.

In the remainder of this section, we present three exhibits that detail the estimates of separate reaction functions for each of the three Chairmen. We estimate the basic equation (equation (4)) using OLS and then re-estimate the equation after eliminating insignificant terms.[16]

13. Still, it is possible that the equilibrium rate was elevated during the Volcker period given the large federal budget deficits. For a model-based definition of a time-varying equilibrium rate, see Bomfim (1997).

14. There is also, of course, the issue of time variation in π^* and r^* (as noted in footnote 13). Even during a given Chairman's term, there may well be changes in the target inflation rate. Indeed, this is the essence of the opportunistic approach to monetary policy described by Bomfim and Rudebusch (1998).

15. This series is conceptually similar and highly correlated with the Q^* series in Braun 1990; Hallman, Porter, and Small 1991; and Orphanides (1997).

16. Given the lags in the transmission process of monetary policy, there is little danger of reverse causation from i_t to π_t and y_t.

Exhibit 1: The Greenspan Period, 1987.Q3–1997.Q4

The lagged gap is insignificant in A, the basic regression, so it is eliminated in B. Regression B explains 71 percent of the quarterly variation in the change in the funds rate (with an adjusted $R^2 = 0.67$), and has a standard error of 27 basis points. Not surprisingly, this regression has a closer fit with the data than Taylor's original specification.[17]

Several interesting issues arise from regression B. First, the estimates suggest gradual, rather than instantaneous, adjustment of the funds rate to the rule. The funds rate typically adjusts enough to eliminate 28 percent of the difference between the lagged actual and rule-recommended funds rate each quarter. Second, the estimated weight on the GDP gap of 0.99 is higher than Taylor assumed (0.50). In this regard, some researchers have found that a larger weight on the output gap than Taylor assumed produces a lower output variance for a given inflation variance in model simulations (e.g., Rudebusch and Svensson 1998 and Williams 1997).

Finally, the data provide a fairly narrow range of estimates of the equilibrium real funds rate and the inflation target. The various estimates of the equilibrium funds rate and the inflation target that are consistent with the estimated constant term can be seen in the Figure in Exhibit 1. The average long-sample and Greenspan-sample real funds rates and the end-of-sample inflation rate are consistent with a fairly tight range of tradeoffs on the line. The estimates of both the inflation target and the real equilibrium funds rate all lie in a range from 1.8 to 2.8 percent—not far from Taylor's assumption of 2 percent.

EXHIBIT 1

REGRESSION RESULTS FOR GREENSPAN PERIOD, 1987.Q3–1997.Q4

	α	γ	λ_1	λ_2	λ_3	ρ	$\bar{R}^2$	*SEE*	*Q*
A	1.21	0.27	0.57	1.10	–0.12	0.43	0.67	0.27	20.32
	(1.79)	(4.87)	(2.72)	(2.83)	(–0.31)	(3.94)			(0.03)
B	1.31	0.28	0.54	0.99		0.42	0.67	0.27	20.78
	(2.26)	(5.27)	(2.96)	(7.46)		(4.00)			(0.02)

GREENSPAN REGRESSION INFORMATION

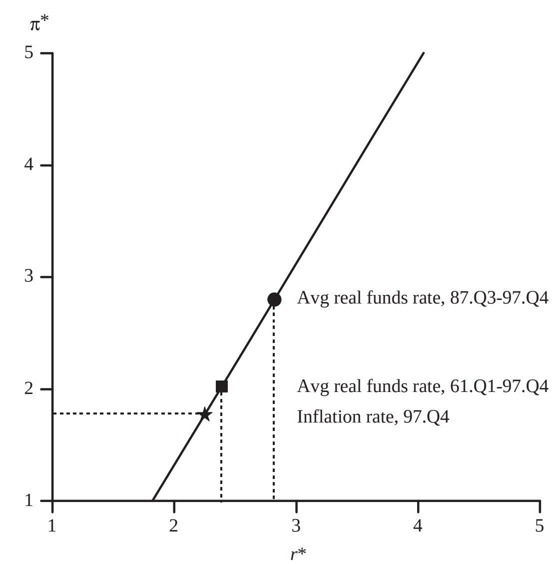


Exhibit 2: The Volcker Period, 1979.Q3–1987.Q2

As with the prior regression, estimation of the basic equation (A) finds evidence of partial adjustment of the funds rate to the rule. However, the dynamic pattern is somewhat different in that the lagged dependent variable is not significant; thus, in regression B, we drop this term. Regression A also suggests that the Volcker period involved a response to the change in, rather than the level of, the GDP gap. Indeed, this restriction cannot be rejected at any conventional significance level. In regression B, the change in the GDP gap enters. The coefficient on the inflation gap in B is very close to the 0.5 assumed by Taylor, although the estimated coefficient is only very marginally statistically significant. Overall, our results suggest that policy was concerned with the rate of inflation relative to a target and with the growth rate of real GDP relative to the growth rate of potential GDP.

However, the equation is estimated with much less precision for the Volcker period than for the Greenspan period. The coefficients on λ_1, λ_2, and λ_3 are individually significant only at the 6 to 8 percent level (although they are jointly significant at the 1 percent level), and the standard error is 1.31 percentage points compared to 0.27 percentage point in the Greenspan period. In part, this could

17. The *Q*-statistic suggests the possibility of autocorrelation in the regression. Much of this may be due to our use of time-aggregated data. When we respecified the regression using end-of-quarter funds rate data, the *Q*-statistic did not show signs of autocorrelation, the lagged change in the funds rate became statistically insignificant, and the other coefficients were close to the results in the original specification. This result adds to our confidence in the specification of the right-hand-side variables in regression B, which we retained in the interest of obtaining an equation that can be used in a quarterly macroeconomic model with quarterly average measurement of the funds rate.

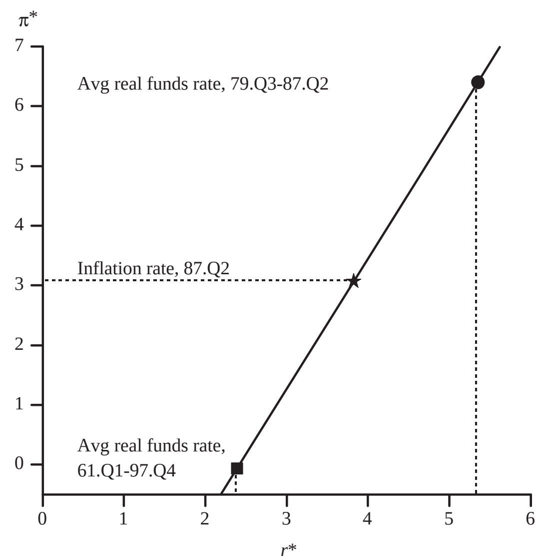
EXHIBIT 2

REGRESSION RESULTS FOR VOLCKER PERIOD, 1979.Q3–1987.Q2

	α	γ	λ_1	λ_2	λ_3	ρ	$\bar{R}^2$	*SEE*	*Q*
A	2.04	0.36	0.69	2.40	–2.04	–0.08	0.47	1.33	10.75
	(0.87)	(2.25)	(1.32)	(1.35)	(–1.42)	(–0.44)			(0.22)
B	2.42	0.44	0.46	1.53*	–1.53*	—	0.48	1.31	9.43
	(1.56)	(3.64)	(1.79)	(1.92)	(–1.92)				(0.31)

* restriction: $\lambda_2 + \lambda_3 = 0$

VOLCKER REGRESSION INFORMATION



be because the policy problem in 1979 was so clear: double-digit inflation then was so far above any reasonable inflation target that policy did not need to be as concerned with the rather refined judgments about funds rate settings provided by a Taylor-style reaction function. Instead, policy could make gross judgments about keeping the real funds rate at a "high" level until inflation began to come down. Alternatively, the imprecision could reflect noisy movements in the funds rate under a nonborrowed reserves operating procedure.

As shown in the Figure, we obtain a wider range of estimates for the implicit inflation target during the Volcker period than during the Greenspan period. This occurs because the average real funds rate during Volcker's tenure (5.35 percent) differs substantially from the average over the entire sample (2.39 percent). The corresponding estimates of the inflation target range from 6.4 percent to –0.1 percent. These estimates bracket the end-of-Volcker-sample inflation rate of 3.07 percent, which corresponds to an r^* of 3.8 percent. The initial tightening of monetary policy could be justified by any of these inflation targets, since inflation was almost 9 percent at the beginning of the Volcker period. Thereafter, it is not possible to tell if the high real funds rates (relative to the Greenspan period) reflects a very low inflation target or a belief that the equilibrium real interest rate was unusually high, possibly because of a perceived need to offset the effects of highly expansionary fiscal policy.

Exhibit 3: The Burns Period, 1970.Q1–1978.Q1

A key feature of this Exhibit is the insignificance of the coefficient on the inflation gap in the general regression A. Note that this does not mean that inflation considerations are entirely absent from the regression for the Burns period. As is clear in equation (2), even when $\lambda_1 = 0$, the nominal funds rate is affected by movements in inflation; however, these movements are simply the one-for-one changes that are necessary to hold the level of the real funds rate unchanged in the face of changes in inflation. Thus, the regressions suggest that the *real* funds rate was *not* adjusted on the basis of changes in inflation.

The lack of a response of the real funds rate to deviations between inflation and an inflation target will be a critical failing for a monetary policy rule. Without the "anchor" of an inflation target to moor the economy, nominal quantities, like inflation and aggregate demand, will be allowed to drift. Indeed, the lack of an implicit inflation target appears to be consistent with the increase in inflation during the Burns period (Figure 1). Of course, other factors may have played a role as well. In particular, there were two large oil shocks in the Burns period. These events no doubt contributed to the inflationary problems of the period, although a consistent policy response to these inflation shocks most likely would have reduced their effects. We address this issue in more detail in the next section.

When the insignificant inflation and contemporaneous gap terms are dropped, we obtain regression B, which shows partial adjustment of the funds rate to a rule that includes only the lagged GDP gap.[18] Since the inflation gap is not in the regression, the constant term (α) provides an estimate of the equilibrium real funds rate (r^*) implicit in Fed

18. This regression shows signs of autocorrelation (the *Q*-statistic has a p-value of 0.6 percent). As with the Greenspan regression, when the

EXHIBIT 3

REGRESSION RESULTS FOR BURNS PERIOD, 1970.Q1–1978.Q1

	α	γ	λ_1	λ_2	λ_3	ρ	$\bar{R}^2$	*SEE*	*Q*
A	1.68	0.56	–0.15	0.16	0.72	0.25	0.53	0.84	15.92
	(1.43)	(4.34)	(–0.80)	(0.45)	(2.59)	(1.67)			(0.04)
	$\alpha = r^*$								
B	0.71	0.58	—	—	0.89	0.26	0.52	0.84	14.81
	(2.68)	(4.78)			(5.85)	(1.76)			(0.06)

policy during the period. (This can be seen in equation (4) by setting $\lambda_1 = 0$, which makes $\alpha = r^*$.) One interpretation of this estimate (i.e., that $r^* = 0.7$ percent) is that policy was predicated on the belief that the equilibrium real funds rate was well below postwar experience in the U.S.

A perhaps more plausible interpretation is that the level of the output gap prevailing at the time was consistently mis-estimated during the Burns period. If, for example, the *average* level of the output gap were estimated to be around $1\frac{1}{2}$ percentage points lower than our current estimate for that period, then the estimates in regression B would be consistent with an *average* equilibrium funds rate of around 2 percent. The existence of such large mistakes in the contemporaneous estimates of the output gap have been given an important role during the period by many analysts (e.g., Blinder 1979, p. 35). Such a consistent string of mistakes would not be too surprising. During the Burns period, productivity and potential output both exhibited a surprising (and still largely unexplained) slowdown in growth, and demographic factors, especially the entrance of the baby boom generation into the labor force, conspired to create an increase in the natural rate of unemployment that also was unexpected. Indeed, during the Burns period, there was a widespread view that an unemployment rate of 4 to 5 percent was a suitable benchmark rate for policy. In contrast, recent (time-varying) estimates of the natural rate that prevailed during the Burns period are in the 6 percent range (e.g., Gordon 1997). Such a difference could account for the consistently easy policy during the Burns period. (With an Okun's Law coefficient of 2, the unemployment gap error translates into an underestimation of the output gap on the order of 2 to 4 percent, which would put the funds rate too low.)

funds rate is defined in terms of the level of the last week in the quarter, rather than as a quarterly average, the coefficients in the equation change very little, but the *Q*-statistic becomes insignificant.

IV. MODEL-BASED EVALUATION OF ALTERNATIVE POLICY RULES

It has become common to evaluate the effectiveness of policy rules or reaction functions like the ones estimated above in terms of the volatility of inflation and output that might result if the rule were used by policymakers. Estimates of this volatility can be obtained from simulations of macro models that include the rule to be evaluated (or by similar analytical methods). See, for example, Rudebusch and Svensson (1998) and Levin, Wieland, and Williams (1997) for recent examples. While exercises of this type can provide useful information for evaluating alternative rules, they are not likely to provide conclusive answers. The results depend upon the particular model employed in the analysis, and because there is no single consensus model in macroeconomics, results from any one model will be subject to debate. (Also, in many cases, the relative rankings of alternative rules are not clear because a tradeoff exists between a rule that has a lower inflation variance and another rule that has a lower real GDP variance.)

As an initial step in evaluating the reaction functions estimated in this paper, we have used a simple model from Rudebusch and Svensson (1998). It includes an aggregate supply equation (or "Phillips curve") that relates inflation to the output gap:

$$(5) \quad \tilde{\pi}_{t+1} = 0.68\tilde{\pi}_t - 0.09\tilde{\pi}_{t-1} + 0.29\tilde{\pi}_{t-2} + 0.12\tilde{\pi}_{t-3} + 0.15\, y_t + \varepsilon_{t+1},$$

(where $\tilde{\pi}_t$ is the quarterly, not four-quarter, inflation rate) and an aggregate demand equation (or "IS curve") that relates output to a short-term interest rate:

$$(6) \quad y_{t+1} = 1.17y_t - 0.27y_{t-1} - 0.09(\bar{i}_t - \pi_t) + \eta_{t+1}$$

(where $\bar{i}_t$ is the average funds rate over the current and prior three quarters). This simple model produces transparent results, captures the spirit of many practical policy-oriented macroeconomic models, and fits the data quite well.[19] In addition to these equations, the estimated reaction functions for the three periods were included one at a time (as well as the original Taylor rule), and the unconditional standard deviations of inflation and the output gap were calculated.

The results are presented in Table 2. In the Rudebusch-Svensson model, the estimated reaction function for the

19. The equations were estimated from 1961.Q1 to 1996.Q4. See Rudebusch and Svensson (1998) for details. The estimates in (5) and (6) differ very slightly from those in that paper because of the longer sample and data revisions.

TABLE 2

MODEL-BASED VOLATILITY RESULTS

MONETARY POLICY REACTION FUNCTION	STANDARD DEVIATION π_t	y_t
Taylor rule	3.86	2.23
Greenspan period	3.87	2.18
Volcker period	4.80	2.73
Burns period	Does not converge	

Greenspan period has an advantage over the function for the Volcker period: the former function produces a lower standard deviation for the real output gap and about the same standard deviation for (four-quarter) inflation. However, we would not want to emphasize this comparison too much because the differences are not large and may be reversed in a different model. The function for the Greenspan period produces about the same volatility of both inflation and real GDP as the original Taylor rule.

The results for the Burns period seem more telling, since the model did not converge when that reaction function was included. This dynamic instability reflects the fact that inflation is not anchored in the Burns period. This result is likely to show up in a variety of models when the reaction function for the Burns period is used. Indeed, Clarida, Gali, and Gertler (1997a) use a calibrated forward-looking model to show that their estimated pre-1979 reaction function is unstable.

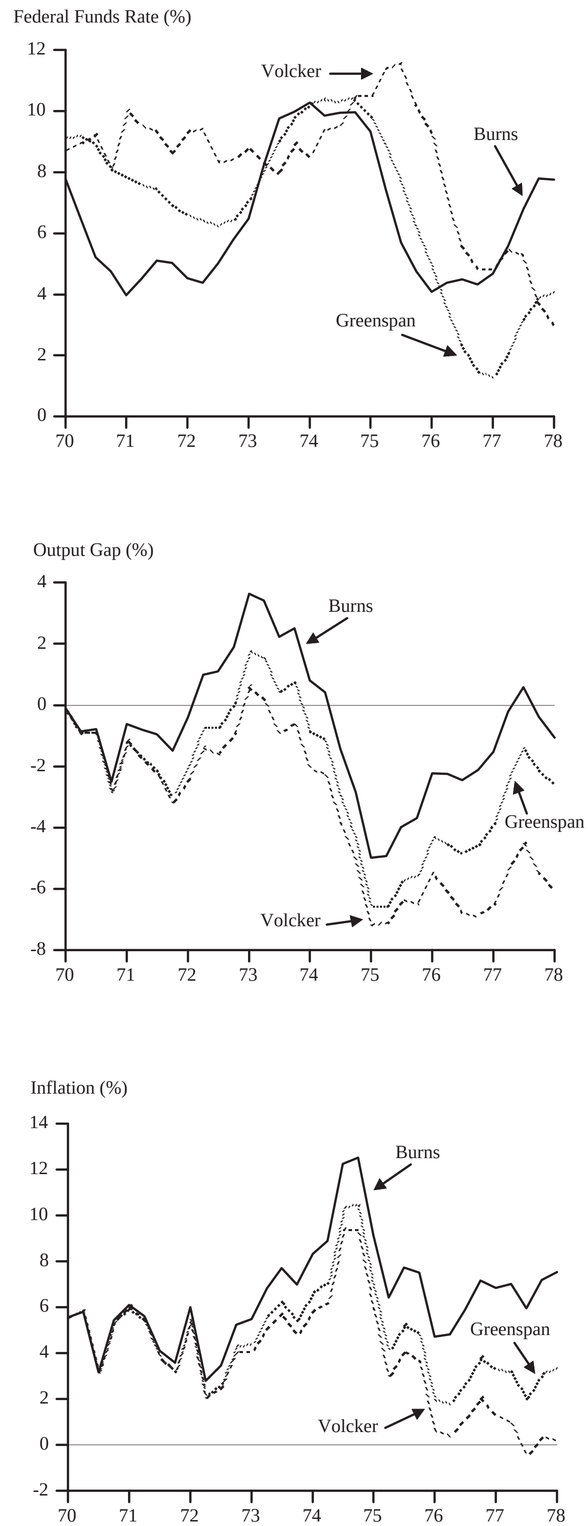
The contrast between the three estimated reaction functions is demonstrated in Figure 2 with counterfactual simulations of the Burns period. These are simulations of equations (5) and (6) along with, in turn, the Burns, Volcker, and Greenspan reaction functions. The actual historical shocks to equations (5) and (6) in the Burns period are used, so in all three cases inflation is pushed up by unfavorable shocks. Still, the difference between the Burns reaction function and the other two is striking, for only with the Burns reaction function does inflation remain at a high level.

V. CONCLUSIONS

The estimates in this paper indicate that a Taylor-type reaction function seems to capture some important elements of monetary policy during Alan Greenspan's tenure to date as Federal Reserve Chairman. This regression implies that movements in the funds rate over that period have been

FIGURE 2

COUNTERFACTUAL SIMULATIONS: 1970.Q1–1978.Q1



broadly consistent with a policy regime aimed at low inflation in the long run and a stable level of output around trend in the short run. However, the results differ somewhat from Taylor's original specification of the rule in two main ways. The funds rate appears to have reacted about twice as strongly to the GDP gap as Taylor assumed, and it appears to have moved gradually, rather than instantaneously, into rough accord with the estimated Taylor rule.

The estimates for the Volcker period are less precise than those for the Greenspan period. Nonetheless, they suggest that the Fed adjusted the funds rate gradually in response to concerns with achieving an inflation target well below the rate inherited by the FOMC in the late 1970s. This result is consistent with the substantial progress achieved in reducing inflation during the period. Policy also appears to have given weight to cyclical considerations, but this concern came in the form of reactions to the growth rate rather than to the level of real GDP.

In the Burns period, we find a weak policy response to inflation. Instead, policy seems to have been geared mainly toward gradual responses to the state of the business cycle. Moreover, some evidence suggests that policy either was oriented around an unusually low estimate of the equilibrium real funds rate or around an estimate of potential output that appears to have been too high in retrospect. These results seem consistent with the key feature of Burns's tenure as Chairman of the Fed—rising inflation—and they appear to show up as dynamic instability in our model simulations.

Overall, the dynamic Taylor-type reaction functions estimated during the Burns, Volcker, and Greenspan periods, appear to have differed in important ways from one another. As noted above, this investigation has not provided a complete representation of changes in Fed behavior, in part because we have controlled for only one source of sample instability. This may account for the sensitivity of some of the results to alternative specifications as shown in the Appendix. However, we hope our results represent a step in the direction of uncovering the key elements—and changes—in Federal Reserve behavior over time.

The finding that the monetary policy regime may have changed in significant ways over time has implications for at least two strands of literature in macroeconomics. First, the finding raises questions about attempts to estimate monetary policy shocks using identified VARs estimated over long sample periods. If the implicit reaction functions in these VARs do not properly capture the changes in the way policy was formulated, then the estimated shocks will not properly measure the "surprises" in policy. Thus, our results reinforce the conclusions of Rudebusch (1998) that such VARs may be misspecified. Second, in macroeconomic models with rational expectations, parameters throughout the models depend upon the monetary policy regime in place. If the policy regime has changed frequently in the postwar period, it may be difficult to obtain good estimates of these rational expectations models, in part because we may not have long enough sample periods under a consistent policy regime.

APPENDIX: ALTERNATIVE SPECIFICATIONS

We examine the robustness of the results presented in the text by looking at regressions using alternative measures of inflation and the GDP gap. The results are presented in Table A1. With regard to inflation, the estimated regressions show little sensitivity to these alternative measures.

With regard to the GDP gap, we estimate reaction functions using three estimates of potential GDP, namely, Y^*, which is described in the text, a segmented linear trend with one break in 1973.Q1, and a quadratic trend. Figure A1 shows the alternative estimates of potential output and the corresponding GDP gaps. The GDP gap measured in terms of Y^* has cross-correlations of 0.99 and 0.80 with the quadratic and linear trend gaps, respectively. A recent example of a divergence among these series occurred in the 1990s, when the segmented linear trend showed output consistently below potential, while the other two measures showed a rising gap that became positive toward the end of the sample. Differences like these can have noticeable effects on Fed policy concerns. The regression results for the reaction function using the linear trend differ from those using the other gap measures in Table A1. In fact, in the Greenspan period, the introduction of the linear trend actually changes the sign of the response to the inflation gap. The alternative measures of the gap have little effect on the results for the Burns or the Volcker periods.

TABLE A1

REACTION FUNCTIONS—ALTERNATIVE SPECIFICATIONS
INFLATION

	α	γ	λ_1	λ_2	λ_3	ρ	r^*	$\bar{R}^2$	*SEE*	*Q*
GREENSPAN										
GDP Deflator	1.31	0.28	0.54	0.99	–	0.42	–	0.67	0.27	20.78
	(2.26)	(5.27)	(2.96)	(7.46)	–	(4.00)	–	–	–	(0.02)
PCE Price Index	2.39	0.23	0.07	1.02	–	0.44	–	0.62	0.29	12.12
	(3.56)	(4.30)	(0.37)	(5.66)	–	(3.89)	–	–	–	(0.28)
Core CPI	1.00	0.25	0.37	1.15	–	0.49	–	0.64	0.28	16.24
	(1.28)	(4.56)	(1.79)	(7.20)	–	(4.47)	–	–	–	(0.09)
VOLCKER										
GDP Deflator	2.42	0.44	0.46	1.53	–1.53	–	–	0.48	1.31	9.43
	(1.56)	(3.64)	(1.79)	(1.92)	(–1.92)	–	–	–	–	(0.31)
PCE Price Index	1.46	0.29	0.54	2.58	–2.58	–	–	0.40	1.41	9.06
	(0.52)	(2.59)	(1.21)	(1.65)	(–1.65)	–	–	–	–	(0.34)
Core CPI	1.32	0.35	0.35	2.57	–2.57	–	–	0.43	1.37	8.70
	(0.57)	(3.03)	(1.11)	(2.07)	(–2.07)	–	–	–	–	(0.37)
BURNS	$\alpha = r^*$									
GDP Deflator	0.71	0.58	–	–	0.89	0.26	0.71	0.52	0.84	14.81
	(2.68)	(4.78)	–	–	(5.85)	(1.76)	(2.68)	–	–	(0.06)
PCE Price Index	0.95	0.51	–	–	0.86	0.17	0.95	0.48	0.87	16.97
	(2.98)	(4.33)	–	–	(4.67)	(1.14)	(2.98)	–	–	(0.03)
Core CPI	1.30	0.56	–	–	1.14	0.34	1.30	0.40	0.94	10.72
	(4.17)	(3.46)	–	–	(6.12)	(1.99)	(4.17)	–	–	(0.22)

REACTION FUNCTIONS—ALTERNATIVE SPECIFICATIONS
POTENTIAL GDP

	α	γ	λ_1	λ_2	λ_3	ρ	r^*	$\bar{R}^2$	*SEE*	*Q*
GREENSPAN										
Y^*	1.31	0.28	0.54	0.99	–	0.42	–	0.67	0.27	20.78
	(2.26)	(5.27)	(2.96)	(7.46)	–	(4.00)	–	–	–	(0.02)
Segmented Linear Trend	5.35	0.18	–0.99	0.90	–	0.54	–	0.58	0.31	10.71
	(3.52)	(3.56)	(–1.93)	(3.99)	–	(4.88)	–	–	–	(0.38)
Quadratic Trend	1.09	0.28	0.37	0.82	–	0.52	–	0.64	0.28	19.10
	(1.80)	(4.77)	(1.90)	(7.05)	–	(4.94)	–	–	–	(0.04)

(continued)

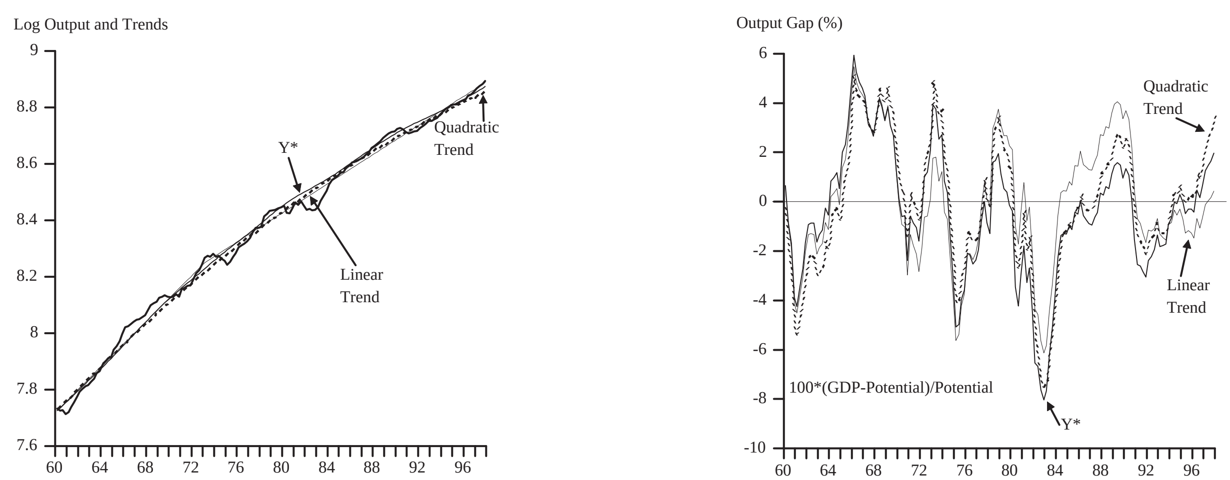
TABLE A1 (Continued)

REACTION FUNCTIONS—ALTERNATIVE SPECIFICATIONS
POTENTIAL GDP

	α	γ	λ_1	λ_2	λ_3	ρ	r^*	$\bar{R}^2$	*SEE*	*Q*
VOLCKER										
*Y**	2.42	0.44	0.46	1.53	–1.53	–	–	0.48	1.31	9.43
	(1.56)	(3.64)	(1.79)	(1.92)	(–1.92)	–	–	–	–	(0.31)
Segmented Linear Trend	2.29	0.43	0.48	1.57	–1.57	–	–	0.48	1.31	9.28
	(1.39)	(3.46)	(1.75)	(1.84)	(–1.84)	–	–	–	–	(0.32)
Quadratic Trend	2.23	0.43	0.50	1.60	–1.60	–	–	0.48	1.31	9.23
	(1.33)	(3.46)	(1.78)	(1.84)	(–1.84)	–	–	–	–	(0.32)
BURNS	$\alpha = r^*$									
*Y**	0.71	0.58	–	–	0.89	0.26	0.71	0.52	0.84	14.81
	(2.68)	(4.78)	–	–	(5.85)	(1.76)	(2.68)	–	–	(0.06)
Segmented Linear Trend	1.54	0.52	–	–	1.07	0.25	1.54	0.46	0.89	13.88
	(3.76)	(4.15)	–	–	(4.83)	(1.59)	(3.76)	–	–	(0.09)
Quadratic Trend	–0.15	0.59	–	–	0.88	0.26	–0.15	0.52	0.84	14.57
	(–0.58)	(4.78)	–	–	(5.90)	(1.78)	(–0.58)	–	–	(0.07)

FIGURE A1

ALTERNATIVE ESTIMATES



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Borrowing Constraints and Asset Market Dynamics: Evidence from the Pacific Basin

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The author would like to thank Mark Spiegel for helpful comments and Priya Ghosh for expert research assistance.

This paper estimates a linearized, stochastic version of Kiyotaki and Moore's (1997) credit cycle model, using land price data from Hong Kong, Japan, and Korea. It is shown that the welfare costs of borrowing constraints are positively related to the persistence of (detrended) land price fluctuations. When the residual demand curve for land is inelastic and the steady state share of land held by the constrained sector is less than 30 percent, welfare costs are less than 1 percent of GDP in all countries. However, the costs of borrowing constraints rise quickly as the constrained sector becomes more important and as the elasticity of unconstrained land demand increases. For example, if the efficient share of the constrained sector is 50 percent and the residual demand elasticity is 2.0, then costs range from 9 percent of GDP in Korea, where fluctuations are relatively transitory, to 11 percent of GDP in Japan, where land price fluctuations are the most persistent.

What is perhaps most surprising about recent events in Asia is not the widespread currency devaluations, but the subsequent declines in economic activity. Many observers noted that the declining yen and the devalued yuan had eroded the competitiveness of these countries. (Chinn 1998 provides some evidence that most of these currencies were overvalued based on standard PPP considerations.) In addition, given these countries' common interest in exporting similar products to the U.S. and Japan, it is not surprising that they devalued together (Huh and Kasa 1997). However, devaluation was supposed to restore their competitiveness and *stimulate* their economies. Instead, these devaluations produced recessions.

There are many reasons why a devaluation might produce a recession.[1] The thesis of this paper is that in the case of the recent Asian crisis, financial market imperfections are a particularly likely explanation. Specifically, I argue that a combination of an open-economy version of Irving Fisher's (1933) debt-deflation hypothesis, featuring foreign debt and a currency devaluation rather than a price level decline as the initial negative impulse, along with leverage-induced feedback between collateralized asset prices, borrowing constraints, and investment as the propagation mechanism, can provide a convincing account of recent events in Asia.

To substantiate this claim, I estimate a linearized version of Kiyotaki and Moore's (1997) credit cycle model. This model features two sectors. One sector is subject to borrowing constraints, i.e., investment must be fully backed by the value of collateral. The other sector is unconstrained and acts as a buffer, i.e., it provides an alternative use for the collateralized asset. Kiyotaki and Moore show that shocks emanating in either sector set into motion a dynamic feedback process between asset prices and borrowing constraints. Fundamentally, this feedback arises from the dual nature of assets in this economy. Not only are durable assets, like land, an input to production, but they also provide collateral, and hence affect borrowing constraints. A sudden decline in asset prices lowers the value of collateral, which reduces investment in the constrained sector. Since in equilibrium the marginal product of capital is

1. Krugman and Taylor (1978) outline a number of demand-side stories, while van Wijnbergen (1986) points to potentially adverse supply effects, e.g., a devaluation inceases the prices imported intermediate inputs.

higher in the constrained sector, a reallocation of investment away from the constrained sector reduces aggregate output, which further depresses asset prices.

Economists have long recognized the potential role of leverage as a cyclical propagation mechanism. For example, Veblen (1904), in his own inimitable way, described the process clearly, if not entirely persuasively:

> Funds obtained on credit are applied to extend the business; competing business men bid up the material items of industrial equipment by the use of funds so obtained; the value of the material items employed in industry advances; the aggregate of values employed in a given undertaking increases, with or without a physical increase of the industrial material engaged; but since an advance of credit rests on the collateral as expressed in terms of value an enhanced value of the property affords a basis for a further extension of credit, and so on. . . . The extension of loans on collateral has therefore in the nature of things a cumulative character. This cumulative extension of credit through the enhancement of prices goes on, if otherwise undisturbed, so long as no adverse price phenomenon obtrudes itself with sufficient force to convict this cumulative enhancement of capitalized values of imbecility. (Chapter 5, p. 55)

Of course, these days economists prefer to study economies inhabited by rational actors, not imbeciles, and the contribution of Kiyotaki and Moore (1997) is to show how such a cumulative process can arise in an explicit, quantifiable, and internally consistent model. They also characterize the (local) dynamics of this process.

In principle, any asset that is not highly specialized could play the role of collateral in a Kiyotaki and Moore-type model.[2] When implementing their model empirically, however, one must take a stand on the exact nature of this collateralizable asset. In this paper, I assume "land" plays the role of collateral (as well as being a factor of production). Land is undeniably a widely used source of collateral. Moreover, casual empiricism suggests that land values go through exactly the sort of boom and bust cycles predicted by Kiyotaki and Moore's model. Unfortunately, "land" is as heterogeneous as "capital" and presents the same sort of measurement and aggregation problems. Also, many other kinds of durable assets are used as collateral, and ignoring these could be misleading in a quantitative exercise.

The remainder of the paper is organized as follows. Section I develops an open-economy OLG version of Kiyotaki and Moore's credit cycle model. For reasons of both analytical convenience and empirical plausibility, I assume the economy is "small," and the world interest rate is given. Even with this simplification the model is nonlinear, and the first order of business is to show that under certain reasonable parameter restrictions, the deterministic steady state is characterized by a unique positive land allocation and associated level of aggregate output. I then incorporate stochastic (non-diversifiable) endowment shocks and linearize around this steady state. To a first-order approximation, land prices and aggregate output turn out to follow stationary AR(1) processes. The dynamics of the current account also are characterized. A key result of the model is the fact that the *persistence* of the model's fluctuations increases as the borrowing constraints become more "important," as measured by the steady state relative size of the constrained sector and the elasticity of the residual demand curve for land. I use this relationship later to back out estimates of the welfare cost of borrowing constraints from estimates of the persistence of land price fluctuations.

Section II provides a brief discussion of the data. I focus on three countries (on an individual, case-by-case basis): Hong Kong, Japan, and Korea. Each of these countries has experienced considerable fluctuations in land values. The exact definition of land differs somewhat from country to country. For Japan and Hong Kong I obtain actual transactions-based data on land prices according to use. I employ broad measures that encompass both residential and commercial uses of land. For Korea the data are closer to being a standard housing price index, which of course is a rather noisy indicator of land values.

Section III begins by presenting trend/cycle decompositions of land prices for each country. Cyclical fluctuations are quite persistent in all the countries, with half-lives of between three to six years. Fluctuations are most persistent in Japan and least persistent in Korea. Since the shocks are regarded as unobservable, the amplitudes of the cyclical fluctuations are harder to interpret.[3] It turns out that the standard deviations of the cyclical components range from a relatively modest 4.5 percent in Korea to a relatively volatile 16 percent in Hong Kong.

Next, using estimates of the persistence parameters, I compute the implied welfare cost of borrowing constraints under alternative assumptions about the structure of the economy. Not too surprisingly, if the steady state share of the constrained sector is small, and land demand is inelastic in the unconstrained sector (so that Harberger triangles are small), then borrowing constraints do not cost the economy very much. For example, a combination of inelastic demand and a constrained sector share of less than 30 per-

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2. Shleifer and Vishuy (1992) discuss how the degree of asset specificity affects the feedback between asset prices and borrowing constraints. Their model is static, however.

3. Thus, this paper focuses more on the second half of the debt-deflation/credit-cycle account of the Asian crisis. That is, in this paper I am more interested in the duration and propagation of the crisis than in the initial impulse that started it.

cent always produces welfare cost estimates of less than 1 percent of GDP, even in Japan, where price fluctuations are the most persistent. However, costs increase rapidly as the constrained sector becomes larger and as residual land demand becomes more elastic. If the elasticity of demand is 2.0, then welfare costs rise to about 10 percent of GDP when the share of the constrained sector is 50 percent and approach 40 percent of GDP if the share of the constrained sector is as high as 70 percent.

Section IV of the paper summarizes the main results and offers a few suggestions for future research.

I. THE MODEL

Kiyotaki and Moore (1997) construct several versions of their credit cycle model, differing in complexity and in the particular dynamic mechanisms highlighted. In each version there are two sectors, a constrained sector and an unconstrained sector. Kiyotaki and Moore refer to the constrained sector as "farming" and to the unconstrained sector as "gathering." Farmers and gatherers are distinguished by the technology available to them for producing (perishable) "fruit." Both technologies use land and labor inputs at time t to produce fruit output at time $t+1$, but differ crucially in the nature of their labor inputs. The labor input of gatherers can be guaranteed ahead of time, independently of any debt they might have. In contrast, farmers cannot commit to work. Hart and Moore (1994) refer to this lack of commitment as the "inalienability of human capital." The inalienability of human capital exposes potential lenders to the risk of default, since it is assumed that no fruit is produced without the farmer's labor input. If a farmer's debt becomes sufficiently onerous, it will be in his interest to withdraw his labor and default on his loan. As a result, lenders will require loans to farmers to be backed by collateral. In general, the amount of collateral required depends on the specifics of the bargaining process that follows default. Based on the results of Hart and Moore (1994), Kiyotaki and Moore argue that farmers will capture the entire difference between their debt and the liquidation value of their land, so that lenders will require the full (expected) value of their land as collateral.[4] In other words, a farmer cannot take out a loan for more than the (expected) value of his current land holdings. This constraint makes the equilibrium sequential and is responsible for all the model's dynamics.[5]

In their "baseline" model, Kiyotaki and Moore make three unconventional assumptions that facilitate the analysis. First, they abstract from issues of risk-sharing by assuming that preferences are linear in fruit consumption. Second, to make the equilibrium interesting, they assume farmers and gatherers have different rates of time preference. In particular, farmers are assumed to be less patient than gatherers, so that in equilibrium farmers are borrowers and gatherers are lenders. Third, they impose a technological upper bound on the savings rate of farmers (by assuming that some of their output is nontradeable) and impose parameter restrictions ensuring a corner solution for their savings decisions. Thus, savings dynamics play no role in the baseline Kiyotaki-Moore model.

Even with these unconventional simplifying assumptions, the model is quite complex, and yields potentially rich dynamic interactions between asset prices and aggregate economic activity. However, the baseline model has a couple of unattractive features that Kiyotaki and Moore address in an extended version. First, there is no aggregate investment in the baseline model. The total supply of land is fixed, and dynamics take the form of reallocations of land between farmers and gatherers. Second, leverage ratios are unrealistically high, being equal to the reciprocal of the gross interest rate. Such high leverage ratios then yield implausibly large impulse responses to unanticipated shocks. In addition, the lack of aggregate investment makes these responses rather transitory.

Kiyotaki and Moore remedy these shortcomings by introducing reproducible capital into the model. This capital takes the form of "trees." Fruit is now assumed to grow on trees using land and labor as inputs, and trees are grown by planting fruit. Trees are assumed to be specific assets, and hence are uncollateralizable. This reduces leverage ratios and dampens the economy's response to shocks. By assuming that in any given period only a fraction of the farmers have the opportunity to invest in trees, Kiyotaki and Moore are also able to draw out the economy's response to shocks. Moreover, they show that this extended model can potentially have (stable) complex roots and thus produce cyclical responses to shocks.

A third version of the model is developed in the appendix to their paper, which is designed to show that none of the substantive results from their baseline model depend

4. Because farmers cannot commit to pay dividends either, introducing an equity market would not help them raise capital. However, in some versions of Kiyotaki and Moore's model, there may be an advantage to setting up a rental market in land.

5. There are of course other ways of introducing financial market imperfections. Perhaps the most common approach is to assume asymmetric information. For example, this is the route taken by Bernanke and Gertler (1989), who study similar issues. However, basing debt on the "inalienability of human capital" rather than on moral hazard or adverse selection simplifies matters considerably in dynamic settings. See Gertler (1992) for a multiperiod application of the asymmetric information approach.

on its unconventional preference and technology assumptions. It is this third version of the Kiyotaki-Moore model that I employ in studying land price dynamics in the Pacific Basin. The model features a Blanchard (1985)-style overlapping-generations structure, in which farmers and gatherers each face a constant probability of dying, $1 - \sigma$, where σ is the probability of surviving from one period to the next. Each period, new cohorts of farmers and gatherers are born, each of size $1 - \sigma$, so that by the law of large numbers the economy's total population remains constant at 2. Although the assumption of geometrically distributed lifetimes is demographically unrealistic, it does have the virtue of greatly facilitating aggregation, since marginal propensities to save are independent of age.

In contrast to the baseline model, preferences of farmers and gatherers are now assumed to be identical and concave. Specifically, both farmers and gatherers have the following logarithmic preferences:

$$\max_{\{c_t\}} E_t \sum_{j=0}^{\infty} (\beta\sigma)^j \ln[c_{t+j}]. \tag{1}$$

That is, both farmers and gatherers maximize the expected present discounted value of utility from fruit consumption, c_t, conditional on surviving each successive period. Note that leisure does not enter the utility function, so that labor is supplied inelastically.

At this point it should be noted that Kiyotaki and Moore conduct their analysis entirely within a context of perfect foresight. This is a useful abstraction when solving the model and illustrating its dynamics. To make the model econometrically implementable, however, non-degenerate shocks must be incorporated. These shocks will play the role of regression error terms. I do this by assuming that each new cohort's endowment is stochastic. In particular, I assume that at time t newborn farmers and newborn gatherers each receive a fruit endowment of e_t. The e_t process is assumed to be independent over time, with constant mean value, $\bar{e}$. I also assume that there is no way to diversify the endowment shock, even though its realization will in general have first-order aggregate effects.

Remember that farmers and gatherers are distinguished by their technologies for producing fruit. Following Kiyotaki and Moore, I assume farmers have linear technologies. Thus, if we denote the time t aggregate land holdings of farmers by $K_{f,t}$ and farmers' fruit output at time $t + 1$ by $Y_{f,t+1}$ we have

$$Y_{f,t+1} = aK_{f,t} , \tag{2}$$

where a is the constant marginal product of land in farming. Gatherers, on the other hand, are assumed to produce fruit subject to diminishing returns. In particular, their production function is quadratic:

$$Y_{g,t+1} = G(K_{g,t}) = b_0 K_{g,t} - \frac{b_1}{2} K_{g,t}^2 , \tag{3}$$

where $Y_{g,t+1}$ is the time $t + 1$ fruit output of gatherers, and $K_{g,t}$ is their time t land holdings. As in the baseline Kiyotaki-Moore model, I assume the total land supply is fixed at $\bar{K}$, so that market clearing requires $\bar{K} = K_{f,t} + K_{g,t}$ for all t. Hence, the model's dynamics take the form of reallocations of land between farmers and gatherers. Although this rules out some potentially important dynamics, these reallocations do incorporate a notion of "fire-sale" asset transactions, which seems to be an issue in the current Asian crisis.[6] To guarantee an interior steady state allocation of land I impose the following parameter restrictions:

$$b_0 > a > a\sigma > b_0 - b_1\bar{K}. \tag{4}$$

This says that if gatherers hold all the land, the marginal product of land in farming is greater than in gathering, whereas if farmers hold all the land, then the marginal product of land in gathering is greater.

At the start of each period, exchange takes place in four markets: (i) a spot commodity market in which fruit is bought and sold, (ii) a real estate market in which land is exchanged, (iii) a domestic bond market in which farmers and gatherers borrow and lend amongst themselves, and (iv) an international capital market that absorbs the difference between domestic production and domestic expenditure. Fruit is assumed to be the numeraire, with price normalized to unity. The time t price of a unit of land is denoted q_t, and the (constant) gross world interest rate is R (both expressed in units of fruit).

Farmers and gatherers solve the maximization problem in (1) subject to a sequence of budget constraints. If b_t denotes the time t debt of either a farmer or a gatherer, then these constraints take the following form (in the aggregate):[7]

$$q_t(K_{f,t} - K_{f,t-1}) + Rb_{t-1} + c_t = \sigma aK_{f,t-1} + (1 - \sigma)e_t + b_t \tag{5}$$

for farmers, and

$$q_t(K_{g,t} - K_{g,t-1}) + Rb_{t-1} + c_t = \sigma G(K_{g,t}) + (1 - \sigma)e_t + b_t \tag{6}$$

for gatherers.

The right-hand sides of these constraints are the sources of time t funds, which consist of current fruit production of surviving farmers and gatherers, endowments of the newborn, and issues of new debt. The left-hand sides are the uses of time t funds, given by land purchases, debt repayments, and consumption expenditures.

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6. Note, however, that foreigners cannot buy land.

7. Mortality risk implies that the interest rate on individual loans is R/σ. However, in the aggregate, this risk is fully diversifiable, so the sectoral budget constraints take the forms in (5) and (6).

The key ingredient of the model is a constraint limiting the debt, b_t, of farmers. This constraint arises from their inability to commit to work, along with the assumption that no fruit is produced without labor. Kiyotaki and Moore argue that with perfect foresight no farmer will be able to take out a loan that exceeds the present value of his current land holdings, given that lenders recognize the incentive of farmers to default if debt were to exceed this value. However, when land prices are stochastic, the future value of collateral is unknown, and it is not clear how this uncertainty will affect the required level of collateral.[8] For simplicity, I just assume that farmers are able to borrow up to the *expected* present value of their land, less an additive "risk premium," $\phi > 0$, so that with uncertainty the borrowing constraint takes the form:[9]

$$b_t \le \frac{1}{R} K_{f,t} E_t q_{t+1} - \phi , \tag{7}$$

where R appears rather than R/σ since when a farmer dies his land remains. In general, one would expect the risk premium ϕ to depend on the left tail of the support of the endowment shock process. That is, when the potential for negative shocks increases, greater collateral will be required. However, absent a formal analysis, this is just a conjecture.

I assume that in equilibrium equation (7) is binding for all realizations of e_t. (This implies a restriction on the relationship between the farmer's marginal product of land, a, and the world interest rate, R, which is derived below.) Using (7) at equality to substitute out for b_t in the (aggregate) budget constraint of farmers yields:

$$\begin{aligned} u_t K_{f,t} + c_t &= \sigma a K_{f,t-1} + (1-\sigma) e_t \\ &\quad + \sigma \varepsilon_t K_{f,t-1} + (R-1)\phi , \end{aligned} \tag{8}$$

where

$$u_t \equiv q_t - \frac{1}{R} E_t q_{t+1} \tag{9}$$

is the time t "user cost of capital," or in this case, the required down payment on a fully mortgaged unit of land, and where $\varepsilon_t \equiv q_t - E_{t-1} q_t$. Thus, the term $\varepsilon_t K_{f,t-1}$ represents unanticipated capital gains from holding land. This term is, of course, missing from Kiyotaki and Moore's perfect foresight version of the model.

Solving (1) subject to the budget constraint in (8) yields the following decision rule for farmers' investment expenditure on land:

$$u_t K_{f,t} = \beta\sigma[\sigma a K_{f,t-1} + (1-\sigma) e_t + \sigma \varepsilon_t K_{f,t-1} + (R-1)\phi]. \tag{10}$$

That is, farmers spend a fixed fraction of their time t net worth on land. The remaining fraction, $1 - \beta\sigma$, is spent on consumption.

Equation (10) is one of the two fundamental equations of the model. The second fundamental equation summarizes optimal behavior by gatherers. Since gatherers do not face borrowing constraints, their land purchases are based on a no-arbitrage condition. In particular, gatherers must be indifferent between lending and buying land (or, alternatively, between borrowing and selling land). This will be the case when the following equality holds:

$$\frac{G'(K_{g,t})}{u_t} = \frac{R}{\sigma} . \tag{11}$$

The left-hand side of (11) is the rate of return from buying a unit of land, and the right-hand side is the return from lending. (Remember that a mortality risk premium is charged on individual loans.)

If we use the market-clearing condition, $\bar{K} = K_{g,t} + K_{f,t}$, and the definition of u_t in equation (9), then equations (10) and (11) can be reduced to two equations in the two unknown stochastic processes, q_t and $K_{f,t}$. If (11) is used to substitute out for u_t in (10), we get the following nonlinear stochastic difference equation that determines the equilibrium path of farmers' land holdings:

$$\begin{aligned} \frac{\sigma}{R} G'(\bar{K} - K_{f,t}) K_{f,t} &= \beta\sigma\Big[\sigma a K_{f,t-1} + (1-\sigma) e_t \\ &\quad + \sigma \varepsilon_t K_{f,t-1} + (R-1)\phi\Big]. \end{aligned} \tag{12}$$

The following two propositions summarize the essential properties of this difference equation.

PROPOSITION 1: *There exists a unique positive steady state allocation of land. If the world interest rate satisfies the restriction,* $R\beta > 1$, *and the production function parameters satisfy the restrictions in (4), then in the steady state farmers' land holdings are:*

(13)

$$K_f^* = \frac{[\beta R \sigma a - (b_0 - b_1 \bar{K})] + \sqrt{[\beta R \sigma a - (b_0 - b_1 \bar{K})]^2 + 4 b_1 \beta R [(1-\sigma)\bar{e} + (R-1)\phi]}}{2 b_1} .$$

8. See Hart (1995, pp. 112–115) for a brief discussion of the complications that arise with uncertainty.

9. Lacker (1998) provides a formal (two-period) analysis of optimal borrowing contracts and discusses the circumstances under which collateralized debt supports an informationally constrained Pareto optimum. In some versions of his model a term like ϕ appears, which is a Lagrange multiplier on the constraint that collateral transfers not exceed the borrower's holdings of collateral. Another possibility would be to incorporate a (multiplicative) margin requirement. See Edison, Luangaram, and Miller (1998) for an analysis of this case.

Proof: From (3), $G'(\bar{K} - K_{f,t}) = b_0 - b_1(\bar{K} - K_{f,t})$. Hence, the left-hand side of (12) is quadratic. To solve for the steady state, set $K_{f,t} = K_{f,t-1} = K_f^*$, $e_t = \bar{e}$, and and $\varepsilon_t = \bar{\varepsilon} = 0$. This gives $b_1 K_f^{*2} + [(b_0 - b_1\bar{K}) - \beta R\sigma a]K_f^* - \beta R[(1-\sigma)\bar{e} + (R-1)\phi] = 0$. By inspection, the product of the roots is negative. Therefore, there is always one positive root and one negative root. Given the parameter restrictions, equation (13) is the positive root. o

Linearizing (12) around K_f^* gives us:

Proposition 2: *In the neighborhood of the steady state, farmers' land holdings follow a stationary AR(1) process given by:*

$$K_{f,t} = K_f^0 + \lambda K_{f,t-1} + \frac{\lambda}{\sigma a}\left[(1-\sigma)e_t + \sigma\bar{K}\varepsilon_t\right], \tag{14}$$

where

$$\lambda = \frac{\beta R\sigma a}{\beta R\sigma a + \sqrt{[\beta R\sigma a - (b_0 - b_1\bar{K})]^2 + 4b_1\beta R[(1-\sigma)\bar{e} + (R-1)\phi]}} \tag{15}$$

and

$$K_f^0 = \frac{b_1\bar{K}^2}{b_0 - b_1\bar{K} + 2b_1K_f^*}\ .$$

Proof: Apply Taylor's theorem to (12) and use (13). o

Once we have determined the equilibrium $K_{f,t}$ process, we can use equation (11) to derive the equilibrium q_t process.

Corollary 1: *In the neighborhood of the steady state, land prices are given by the following stationary AR(1) process:*

$$\begin{aligned} q_t &= \bar{q} + \left[\frac{\sigma b_1}{R-\lambda}\right]K_{f,t} \\ &= \left[(1-\lambda)\bar{q} + K_f^*\frac{\sigma b_1}{R-\lambda}\right] + \lambda q_{t-1} \\ &\quad + \frac{\lambda b_1}{a(R-\lambda)}\left[(1-\sigma)e_t + \sigma\bar{K}\varepsilon_t\right] \end{aligned} \tag{16}$$

where

$$\bar{q} = \frac{\sigma(b_0 - b_1\bar{K})}{R-1} - \frac{\sigma b_1 K_f^0}{(R-\lambda)(1-\lambda)}\ .$$

Proof: From (9) and (13) we have $q_t - {}^1\!/_R(E_t q_{t+1}) = {}^\sigma\!/_R[b_0 - b_1(\bar{K} - K_{f,t})]$. Iterating forward (i.e., applying a transversality condition on land prices), plugging in for $K_{f,t}$ from (14), and then evaluating the resulting expected present discounted value gives equation (16). o

Later it will be shown that the parameter λ increases when borrowing constraints become more important. Thus, from inspection of equations (14) and (16), borrowing constraints both magnify and prolong the economy's response to shocks.

So far, the fact that the economy is open and has access to world capital markets has been kept in the background. At this point we need to bring international considerations to the foreground. Since R is exogenous, there is no guarantee that for any given value of R domestic expenditures will equal domestic production. If they don't, then the country will have a current account deficit or surplus. The next order of business, therefore, is to characterize the stochastic properties of the current account and make sure it is well-behaved.

By definition, the current account surplus is equal to the trade surplus plus net interest receipts on foreign assets. Letting F_{t-1} denote the stock of net foreign assets at the end of period $t-1$ we have:

$$CA_t = Y_t - C_t - I_t + (R-1)F_{t-1}\,. \tag{17}$$

Differencing both sides and using the identity $CA_t = F_t - F_{t-1}$ give us:

$$CA_t = R\cdot CA_{t-1} + \Delta Y_t - \Delta C_t - \Delta I_t\,. \tag{18}$$

The first thing to note is that in this economy $I_t = 0$ for all t. This is simply because the aggregate supply of land is fixed. Land changes hands, but there is no way to augment its supply. Hence, in the aggregate, investment is always zero. The second thing to note is that as long as $R < 1/\beta\sigma$ farmers will not want to lend. To see this, note from (10) that if ϕ and realizations of e_t, are "small" relative to $K_{f,t}$, then the steady state value of u_t is approximately

$$u^* \approx \beta a\sigma^2\,.$$

Therefore, since the steady state rate of return on land in farming is a/u^* and the rate of return on lending is R/σ, if $R < 1/\beta\sigma$, farming is more attractive than lending.

As a result, since farmers cannot *borrow* on the international capital market, only the decisions of gatherers have a direct bearing on the current account. Given log preferences, gatherers allocate their net worth between investment and consumption in the same way that farmers do. However since gatherers do not face borrowing constraints, their net worth is given by a conventional present value calculation, and their consumption decisions resemble the standard permanent income hypothesis.

A gatherer's net worth has two components. First, unlike farmers, gatherers have a concave technology that yields a stream of profits, denoted by π_t. These profits are the difference between revenue from selling next period's fruit output, discounted at the individual interest rate R/σ, and the current cost of land inputs. That is,

$$\pi_t = \frac{\sigma}{R} G(K_{g,t}) - u_t K_{g,t} . \tag{19}$$

The capitalized value of these profits, denoted Π_t, is given by:

$$\Pi_t = E_t \sum_{j=0}^{\infty} \left(\frac{\sigma}{R}\right)^j \pi_{t+j} . \tag{20}$$

The second component of gatherers' aggregate net worth is their pre-existing holdings of foreign assets, F_{t-1} (or foreign debt if negative). Combining these two components, the aggregate net worth of gatherers, V_t, is:

$$V_t = \Pi_t + R \cdot F_{t-1} . \tag{21}$$

Aggregate consumption can then be written as,

$$C_t = (1 - \beta\sigma)\,(Y_t + q_t \bar{K} + V_t). \tag{22}$$

That is, aggregate consumption is just a fixed fraction of the economy's net worth, where net worth is given by the flow of current fruit output, the value of land, the present value of gatherers' profits, and net foreign assets.

We are now in a position to characterize the equilibrium dynamics of the current account.

PROPOSITION 3: *In the neighborhood of the steady state the economy's current account is a stationary ARMA(2,1) process, which has the following representation:*

$$\begin{aligned} CA_t &= \beta\sigma CA_{t-1} + \beta\sigma\Delta Y_t \\ &\quad - (1-\beta\sigma)\left[\bar{K}\Delta q_t + E_{t-1}\sum_{j=0}^{\infty}\left(\frac{\sigma}{R}\right)^j \Delta\pi_{t+j} + \omega_t \right], \end{aligned} \tag{23}$$

where the i.i.d. process ω_t represents revisions in expectations of future profits and, hence, is uncorrelated with the time-$(t-1)$ information set.

PROOF: Plug (22) into (18) using the definitions in (20) and (21), and use the fact that $\Delta F_{t-1} = CA_{t-1}$. To verify that (23) is ARMA(2,1), evaluate the present value, use (16) to substitute for q_t, apply the operator $(1 - \lambda L)$ to both sides, and then use Granger's Lemma to express the numerator polynomial as an MA(1). o

In sum, it has been shown that this economy features a unique positive steady state allocation of land, that in the neighborhood of this steady state land prices and the equilibrium allocation of land follow stationary AR(1) processes (with identical AR roots), and finally, that the economy's current account balance follows a stationary ARMA(2,1) process, with AR roots $\beta\sigma$ and λ.

Our final task is to investigate the welfare economics of this equilibrium. The essential aspect of the equilibrium from a welfare standpoint is that as long as $R < 1/\beta\sigma$, too little land is held by farmers in the steady state. To see this, remember that the steady state rate of return on farming is a/u^*. Since gatherers are free to equate margins, their steady state rate of return is the market rate, i.e., R/σ. Plugging in the previous (approximate) expression for u^* shows that in the steady state the rate of return in farming exceeds the rate of return in gathering. This discrepancy is a crucial feature of the model. It implies that marginal reallocations of land have first-order consequences for output and asset prices.

In reckoning the welfare cost of borrowing constraints I will follow standard practice and compute the area of a Harberger triangle, which is implied by this return differential. To make the result free of units, I express the cost as a share of GDP. The result is given by the following proposition.

PROPOSITION 4: *The steady state welfare cost of borrowing constraints, expressed as a share of GDP, is given by:*

$$\frac{WC}{Y} = \left(\frac{1-s}{8\eta}\right)\left[\left(1 - \frac{\beta R\sigma}{\lambda}\right) + \frac{s}{1-s}\eta\right]^2 , \tag{24}$$

where η is the absolute value of the elasticity of gatherers' land demand, evaluated at the first-best equilibrium, and s is the first-best equilibrium share of land allocated to farmers.

PROOF: First, the area of the Harberger triangle is $\frac{1}{2}[a - G'(K_g^*)][K_f^{\text{opt}} - K_f^*]$, where K_f^{opt} denotes the first-best allocation of land to farmers. Besides this, every element of the triangle has already been computed. To compute K_f^{opt} we just need to find the value of K that equates the marginal product of land in farming to the marginal product of land in gathering. By direct calculation, this turns out to be:

$$K_f^{\text{opt}} = \frac{a - (b_0 - b_1\bar{K})}{b_1} .$$

Using this with (13) gives us

$$K_f^{\text{opt}} - K_f^* = \frac{2a - (b_0 - b_1\bar{K}) - \beta R\sigma a/\lambda}{2b_1} , \tag{25}$$

where use has been made of equation (15). The rest follows from straightforward algebraic manipulation, noting that in the first-best equilibrium $Y = a\underline{K}$. o

As you would expect, the welfare costs of borrowing constraints increase as the elasticity of demand increases and as the (first-best) steady state share of the farming sector increases. More interesting is the following result, which shows that welfare costs increase with the persistence of land price (and output) fluctuations.

Corollary 2: *If* $\lambda > \beta R\sigma(1 - s)/(1 - s + s\eta)$, *then the steady state welfare costs of borrowing constraints are increasing with the persistence of land price (and output) fluctuations.*

Proof: Differentiate (24) with respect to λ, and verify that $\partial(WC)/\partial\lambda > 0$ if $\lambda > \beta R\sigma(1 - s)/(1 - s + s\eta)$. o

The condition on λ in Corollary 2 derives from the fact that even in the first-best equilibrium there will be some persistence in the economy's response to shocks. As it turns out, $\beta R\sigma(1 - s)/(1 - s + s\eta)$ is the first-best equilibrium value of λ. Notice that while increases in s and η raise the cost of borrowing constraints, they reduce the first-best equilibrium value of λ.

In Section III I use equation (24), along with econometric estimates of λ, to compute estimates of the welfare costs of borrowing constraints for a set of Pacific Basin countries. Before we get to these results, however, I first discuss the data.

II. The Data

In principle, when estimating and evaluating this model one would like to obtain broad measures of collateral, including equipment, structures, land, etc., as well as actual transactions-based data on prices. Also, since it is dynamics that we are interested in, it would be desirable to obtain long enough time series to permit accurate estimates of auto- and cross-correlations. Unfortunately, in practice these data are rarely available, particularly for developing countries. As a result, I limit the scope of the analysis to just three countries—Hong Kong, Japan, and Korea. Moreover, I ignore all forms of collateral other than "land."

The data from these countries are of varying quality. Hong Kong and Japan publish data on actual transactions-based land prices, disaggregated according to use. For both countries I use a series that consolidates the residential and commercial sectors. The data for Hong Kong are from Table 5.9 in the *Hong Kong Monthly Digest of Statistics* (various issues). The series is labeled "Private Domestic/Overall" Although this series is available on a quarterly basis, I sample at an annual frequency. The data run from 1976 through 1997. The data for Japan are from Table 119 in the Bank of Japan's *Economic Statistics Annual* (various issues). The series is labeled "Land Price Indexes of Urban Districts/All Urban Districts."

It is based on surveys of residential, industrial, and commercial land prices in 223 Japanese cities. Although the series is available on a semiannual basis, I again sample at an annual frequency. The data run from 1957 through 1997.

The data for Korea are of more dubious quality. Rather than being observations on land *prices*, they are just indices of the cost of *housing*. Changes in this kind of index more likely reflect changes in rental rates than in capitalized values. Of course, rental rates and purchase prices should be highly correlated, but changes in interest rates could weaken the link.

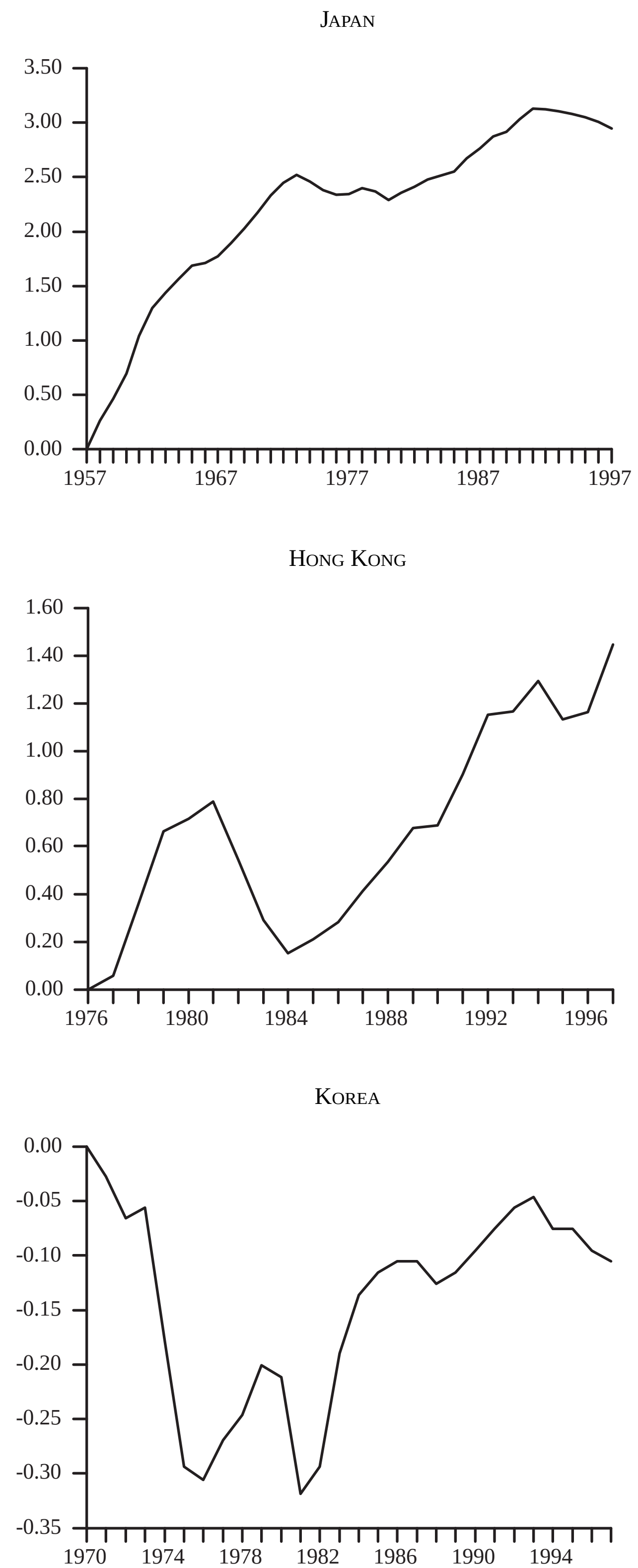
The index for Korea is from Table 103 in the Bank of Korea's *Economic Statistics Yearbook* (various issues). This table provides data on the "All Cities Consumer Price Indexes," and I use the series corresponding to the "Housing Rent" subgroup. The data are sampled annually from 1970 through 1997.

Figure 1 contains plots of each "land price" series. Each series has been deflated by the overall CPI and is graphed on a logarithmic scale. The plots for Japan and Hong Kong seem broadly consistent with informal verbal accounts of their real estate markets. According to these data, Japanese land prices peaked in 1991, and since then have fallen on average by about 18%. Of course, certain segments of the market have declined much more than this (e.g., prime commercial space in downtown Tokyo), but given the rather inclusive definition of the series, an 18% drop seems about right. Notice that an even greater decline occurred in Hong Kong's real estate market during the early 1980s. This of course reflected uncertainty associated with the Sino-British negotiations that were taking place at the time, which also triggered declines in the foreign exchange and stock markets. Since the data end in 1997, the significant declines that occurred in Hong Kong during 1998 as a result of the Asian crisis do not show up here. In fact, the plot reveals that until the crisis hit, the Hong Kong real estate market had been experiencing a boom.

Turning to Korea, the feature that stands out is the dramatic fall in "land prices" that took place during the early 1970s. According to the figure, real land prices declined by over 30% from 1970 to 1975. However, most of this is due to the 1973–74 oil shock, to which Korea was especially vulnerable. During 1973 and 1974 Korean inflation averaged about 25%, so part of the decline probably reflects more of a terms of trade shock than anything else. In fact, in absolute terms land prices rose during the period. The other thing that stands out is that the real estate market in Korea appears to have suffered for several years before the crisis hit at the end of 1997. According to the figure, land prices actually peaked in 1993.

FIGURE 1

LOG OF RELATIVE LAND PRICE



III. EMPIRICAL RESULTS

Kiyotaki and Moore construct their model for the express purpose of studying fluctuations. In doing this, it is useful to abstract from growth. However, the first thing that confronts you when taking the model to data is the presence of trends in land prices (and the prices of other collateralizable assets, for that matter). It would of course be preferable to model trends and cycles simultaneously. One thing we've learned from the Real Business Cycle literature is that factors causing growth can have important cyclical consequences. Nevertheless, the model is complicated enough already, and for now at least I handle trends in the time-honored manner of just mechanically detrending by regressing the logarithms of the series on a linear time trend, with due acknowledgement to the work of Nelson and Kang (1981) on the dangers of inducing spurious cyclicality as a result.[10]

Figure 2 presents the detrended land price series. Along with each series I also plot the fitted values from an AR(1), which according to the model, should characterize the cyclical component of land prices. Not surprisingly, the AR coefficients are highly significant, and imply a high degree of persistence. Estimates range from 0.764 in Korea to 0.867 in Japan. Hong Kong lies in the middle, with a λ estimate of 0.806. These estimates imply that land price cycles have half-lives of between three and five years.

Two notes of caution should be raised about these estimates. First, it is apparent that substantial autocorrelation remains after fitting an AR(1) to detrended land prices. In each case, a second lag enters significantly. Interestingly, estimates from an AR(2) imply humpshaped impulse responses, in which shocks at first cumulate for a few years, as opposed to the monotonic AR(1) dynamics of the model. Second, from Nelson and Kang (1981) we know that fitting a linear time trend to a random walk produces on average a first-order autocorrelation in the residuals of about $1 - 10/T$, where T is the sample size. Given a 25- to 35-year sample we would expect to obtain λ estimates of between 0.6 and 0.7, even when the true data-generating process contains no cyclical component.

With these caveats in mind, Figure 3 uses the estimates of λ to plot out equation (24) for each country. This gives us a measure of the "welfare" or efficiency costs of borrowing constraints, expressed as a share of GDP. Doing this, however, first requires the specification of several free parameters. A reasonable value for the product, $\beta R\sigma$, is

10. I have done some experimenting with univariate Beveridge-Nelson decompositions. Estimates of the cyclical component of land prices turn out to be qualitatively similar, but overall, somewhat less persistent.

FIGURE 2

ACTUAL VS. FITTED CYCLICAL COMPONENTS

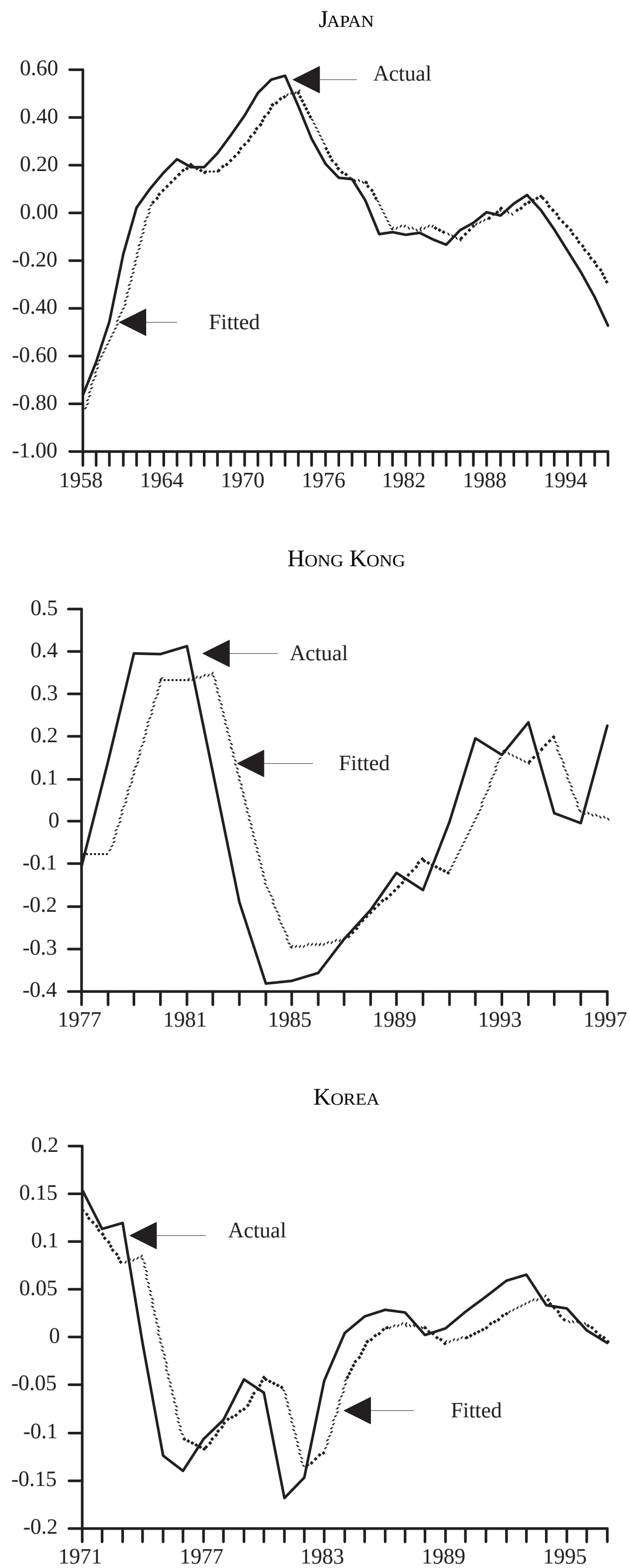
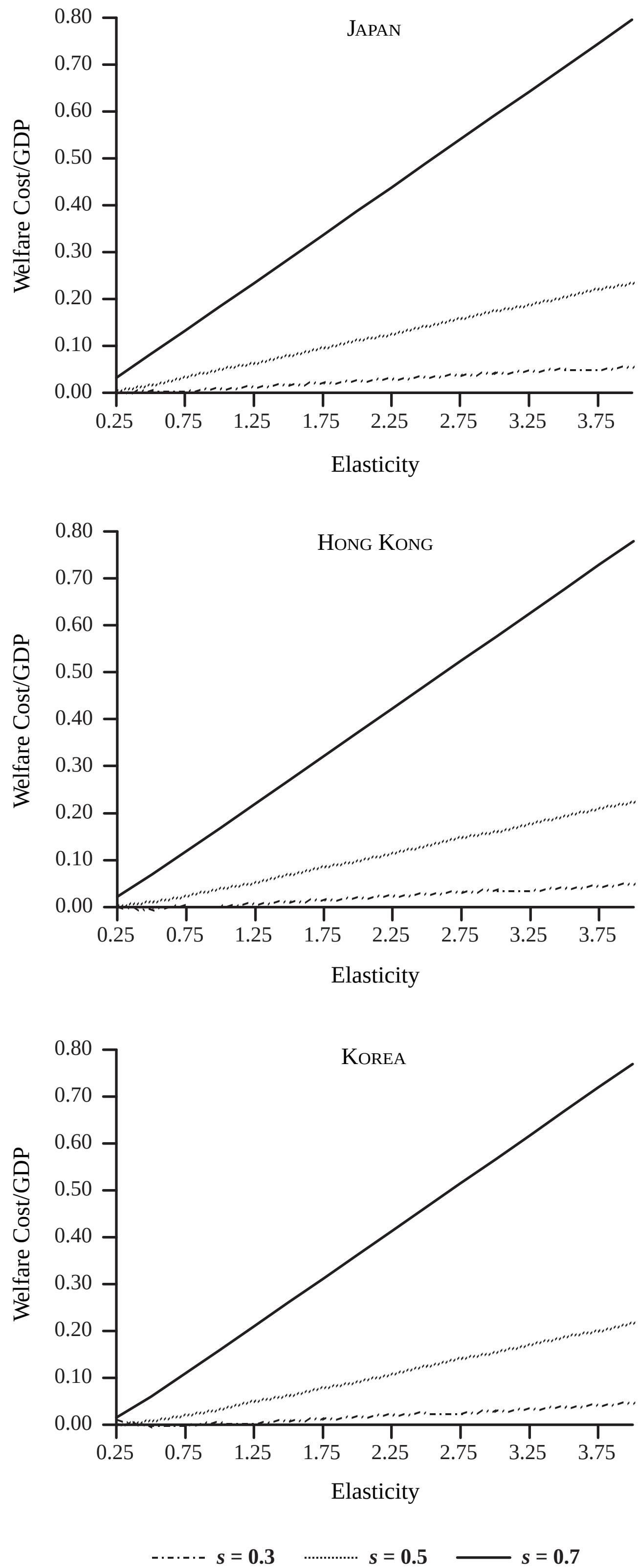


FIGURE 3

ESTIMATES OF WELFARE COSTS



relatively easy to obtain. The model restricts the world interest rate to lie between $1/\beta\sigma$ and $1/\beta$. Economically plausible values of β and σ exceed 0.95, so this is a relatively tight range. Without much loss in generality, I just split the difference and assume R lies at the midpoint of this range, so that $R = 0.5(1/\beta\sigma + 1/\beta)$, which then implies $\beta R\sigma = 0.5(1 + \sigma)$. Thus, we are left with just specifying the demographic parameter, σ. I assume $\sigma = 0.96$, which implies a 25-year planning horizon. The results are insensitive to small perturbations of this parameter.

The remaining two parameters, s and η, are more difficult to specify a priori. These parameters measure the size of the constrained sector and the ease with which land can be transferred between farming and gathering. Since these structural features of the economy are likely to be country-specific and are unidentified in any case given our single parameter estimate, I just plot out the welfare cost function for a grid of η values for each of three different s values. The η values range from a highly inelastic value of 0.25 to a relatively elastic value of 4.0. The share parameter takes on values of 0.3, 0.5, and 0.7.

Evidently, if the share of the constrained sector is less than 0.3, borrowing constraints do not cost the economy much forgone output, regardless of the elasticity of demand. Welfare cost estimates never exceed 6 percent of GDP, even for the highest values of η and λ. However, it is clear that costs rise more than proportionately with s. By the time s reaches 0.7, borrowing constraints are consuming 25−40 percent of the economy's output for intermediate values of η.

Looking across countries, we know that given our λ estimates, Japan will have the highest cost of borrowing constraints and Korea will have the lowest (all else equal). However, the costs are not that sensitive to variations in λ. For example, if $s = 0.5$ and $\eta = 2$, costs range from 9 percent of GDP in Korea, where $\lambda = 0.764$, to 11 percent of GDP in Japan, where $\lambda = 0.867$. Thus, if we assume that roughly half the firms and households in these countries face binding borrowing constraints, then given the size of their economies relative to the U.S., where annual per capita income is about \$25,000, we can say that borrowing constraints cost each person about \$1,667 per year in Japan and Hong Kong, where per capita GDP is roughly two-thirds of U.S. per capita GDP, while they cost about \$1,012 per year in Korea, where per capita GDP is about 45 percent of U.S. per capita GDP.

IV. CONCLUSION

This paper has applied a version of the Kiyotaki-Moore credit cycle model to land price data in Hong Kong, Japan, and Korea. It was shown that land prices can be approximated by an AR(1) process, where the AR coefficient depends positively on the importance of borrowing constraints. It was also shown that borrowing constraints accentuate the economy's initial response to shocks. From a welfare standpoint, it was shown that inferences about the efficiency costs of borrowing constraints can be drawn from estimates of the persistence of land price fluctuations. All else equal, greater persistence implies larger costs. It turns out that estimates of welfare costs are quite sensitive to the steady state share of the constrained sector, which is a parameter that is left unidentified by the model. Based on the parameter estimates, the model suggests that if the share of the constrained sector is between 30–50 percent of the economy, then the welfare costs of borrowing constraints are in the range of 1–10 percent of GDP.

Perhaps the most serious shortcoming of this analysis from the perspective of trying to understand the recent "Asian crisis" is its lack of attention to the source and magnitude of the initial negative impulse(s) that initiated the crisis. For the most part, this paper has focused on the *propagation* of shocks. The model demonstrates that leverage effects can greatly prolong an economy's response to shocks, just as Veblen had conjectured nearly 100 years ago. To the extent that a Kiyotaki-Moore model accurately describes the economies of Asia, one could argue that,absent outside intervention, we should not expect the crisis to abate anytime soon.

A promising avenue for future work would be to try to combine the impulse and propagation mechanisms within a single analytical framework. As recent work by Azariadis and Smith (1998) and Edison, Luangaram, and Miller (1998) has shown, these kinds of models are capable of producing dynamics that are much more exotic than stationary autoregressions. For example, Azariadis and Smith show that multiple steady states can arise, which then opens the door to sunspot equilibria that switch between booms and busts. This would be one way to unite the impulse and propagation problems within a single model.

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Changes in the Structure and Duration of U.S. Unemployment, 1967–1998

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I thank Mary Daly for her detailed comments and Fred Furlong and Tim Cogley for their advice on estimation issues, without implicating any of them for mistakes in implementation. Ken Fujii and Louise Tench provided helpful research assistance.

The unemployment rate is determined by the incidence and duration of unemployment spells. In this paper, I examine the time-series properties of unemployment incidence by reason and the duration of a typical unemployment spell. In line with earlier research, I find strong countercyclicality in unemployment durations, which is relatively uniform across the different reasons for entry into unemployment. However, I also uncover an upward trend in duration that is entirely attributable to rising incidence and duration of permanent job loss. These changes in the structure and duration of unemployment have various policy implications.

In 1998, the U.S. economy produced its lowest unemployment rates since the late 1960s. Despite this, the duration of unemployment spells has remained long compared to typical durations during previous expansions. This is consistent with the long-run secular trend toward rising unemployment durations noted by a number of analysts (e.g., Murphy and Topel 1987, Juhn, Murphy, and Topel 1991, and Sider 1985). Moreover, the structure of unemployment by reason during the 1990s expansion has remained heavily weighted towards permanent job loss rather than voluntary job search and labor force entry decisions. In this paper, I examine the time-series properties of unemployment duration and unemployment incidence by reason, and I investigate the links between them.

Although the unemployment rate is our primary labor market indicator, the underlying distribution of unemployment durations has played a prominent role in the macroeconomics and labor economics literatures since the 1970s. Whether unemployment spells are best characterized as long or short provides information about whether unemployment primarily is voluntary or instead reflects persistent insufficiency of aggregate demand. This information in turn provides insights into what, if any, macroeconomic policies might be appropriate to combat unemployment. The time-series properties of unemployment durations also are relevant to macroeconomic and labor market policies. For example, countercyclical variation in unemployment durations has implications for the behavioral effects and financing of unemployment insurance payments. Moreover, underlying secular trends in unemployment duration may imply long-run changes in the degree of labor market slack associated with a given unemployment rate.

A number of authors (notably Perry 1972, Sider 1985, and Baker 1992a) have investigated the cyclical component of time-series variation in unemployment duration. In this paper, I update earlier results regarding the cyclical sensitivity of unemployment durations, and I extend the analysis by examining secular trends in expected duration and incidence in more detail than did past work. A key component of this formulation involves linking reasons for the incidence of unemployment with changes in duration. In particular, the rising incidence of permanent job loss may be linked to a secular trend toward rising durations, because permanent job losers on average endure substantially longer spells of unemployment than do individuals unemployed for other reasons.

In Section I, I discuss previous work on the time-series properties of unemployment duration and incidence. I describe the data that I use in Section II; these data come from the U.S. Bureau of Labor Statistics' (BLS) monthly household survey. My sample period begins in 1967 and extends to May 1998. Because these data contain information on spells in progress, their use requires estimation of expected completed duration, as described in Section II. Section III presents estimation results; these include basic tabulations for the measures of unemployment incidence and expected duration and regression models that estimate their cyclical and secular properties. I summarize the results in the concluding section, where I also discuss some implications for macroeconomic performance and policy.

I. Unemployment Duration and Incidence

The unemployment rate reflects both unemployment incidence and duration. Although the unemployment rate by itself is our key indicator of labor market conditions, the underlying distribution of unemployment spell durations provides important additional information. In the 1970s and 1980s, the characterization of unemployment durations as "long" or "short" was the subject of substantial debate. The short view, as described most comprehensively by Feldstein (1973), emphasized the dynamic nature of unemployment. Proponents of the short view focused on job turnover and unemployment flows and argued that the pool of unemployed on average is characterized by a large number of individuals who experience relatively short spells of unemployment (i.e., a month or two). This view generally is consistent with voluntary search activity by unemployed individuals or implicit agreements between workers and firms regarding the use of temporary layoffs. In contrast, advocates of the long view, such as Clark and Summers (1979) and Akerlof and Main (1980, 1981), argued that the pool of unemployed typically is dominated by a small number of individuals who experience relatively long spells of unemployment, and who are best described as "involuntarily" unemployed.

These alternative views played an important role in the debate over appropriate macroeconomic and manpower policies aimed at combatting unemployment. In general, the short view is consistent with a less activist policy, given its implications concerning the voluntary nature of unemployment, widespread sharing of the burden of unemployment, and the implied efficiency of the associated employment and unemployment flows. In contrast, the long view of unemployment argues for more activist economic policy, under the assumption that persistent lengthy unemployment spells reflect a shortage of available jobs rather than an equilibrium matching process with frictions.[1]

A related issue is variation in unemployment durations over the business cycle. Changes in the duration distribution of unemployment spells over the business cycle provide information about the degree of demand insufficiency during cyclical downturns and its implications for the matching process between workers and firms. Moreover, changes in the incidence and duration of unemployment over the business cycle may have implications for optimal unemployment insurance policies. For example, Katz and Meyer (1990) documented the importance of temporary layoffs during the early 1980s, and they discuss the distortionary effects of unemployment insurance (UI) benefits on firms' and workers' behavior regarding temporary layoffs.

Kaitz (1970) and Perry (1972) made key early contributions to the analysis of the cyclical properties of unemployment duration. Kaitz used data for the years 1948–1969 and found greater cyclical variability in duration than did Perry for the years 1954–1971. These differences are explained by various differences in approach. Perry also found an upward time trend in duration during the period covered. Sider (1985) updated Kaitz's and Perry's work, emphasizing the importance of using nonsteady-state measures of unemployment duration when cyclical variability is of primary interest. Sider found evidence of substantial cyclical variability, along with an upward time trend for the period 1968–1982.

Michael Baker (1992a) updated these estimates to the late 1980s. He made a key contribution by noting that in contrast to analyses based on aggregate data, which typically uncovered countercyclicality in unemployment durations, some analyses that relied on individual data uncovered procyclicality. One explanation for this discrepancy is the influence of heterogeneity in unemployment incidence: if individuals or groups with long expected duration are more likely to enter unemployment during a downturn, aggregate data will display countercyclicality in unemployment durations even if expected unemployment duration for individuals displays no cyclical properties. Baker applied a decomposition technique that enabled a direct test of the heterogeneity hypothesis. The results did not support the heterogeneity interpretation. In particular, Baker found that pronounced countercyclical variation in duration during the 1980s was attributable to relatively uniform countercycli-

1. Of course, if the long-term unemployed simply lack the appropriate skills to acquire available jobs, aggregate demand management policies are likely to be ineffective.

cality across groups, rather than increasing incidence during downturns for groups with lengthy expected durations.

Although these papers also investigated the cyclical properties of unemployment incidence, none focused directly on the links between secular trends in duration and incidence. I focus on such links, particularly the link between secular increases in unemployment durations and the incidence of permanent job loss. Given that permanent job losers suffer longer spells of unemployment than do individuals unemployed for other reasons, it is likely that an increasing incidence of permanent job loss has contributed to the trend toward rising durations. As discussed in the conclusion, such a secular trend in aggregate labor market outcomes has potentially important implications for macroeconomic policy.

II. ESTIMATING UNEMPLOYMENT INCIDENCE AND EXPECTED COMPLETED DURATION

CPS Unemployment Data

I use monthly data on unemployment levels and rates, which are published by the BLS; the data are available beginning in 1948 and extending to the most recent survey month (May 1998 at the time this paper was written). The data reflect population-weighted counts from the Current Population Survey (CPS), the monthly household survey upon which official labor force statistics are based. The underlying sample is the civilian population aged 16 and over. The published BLS data include information on the total number unemployed, the number unemployed by reason, and the duration of spells in progress. I restrict my analysis to the period beginning in 1967, due primarily to relative consistency in the CPS survey since then. Some analyses are restricted to begin in 1976, because data on unemployment by reason only became available beginning in that year.

The reasons for unemployment identified in the survey fall into five categories: job losers, for whom the survey distinguishes between those on temporary layoff (i.e., they expect recall to the firm from which they were laid off) and permanent job losers (permanent layoffs, firings, or completion of temporary jobs); voluntary job leavers; re-entrants to the labor force; and new entrants to the labor force. For total unemployment and unemployment by reason, the BLS data provide information on the monthly inflow into unemployment. As described in the next subsection, this is the key information used to form a steady-state estimate of expected completed duration.[2]

Use of the BLS unemployment data in a time-series framework requires that an adjustment be applied to the data beginning in January 1994, due to a significant redesign of the survey that became effective at that time; the labor force questions otherwise had been largely unchanged since 1967. As described in Cohany, Polivka, and Rothgeb (1994), results from a parallel survey administered in 1993 indicate that the new survey instrument produces lengthier unemployment durations and changes in unemployment shares by reason. The largest changes in shares are a substantial increase in re-entrants and corresponding decline in new entrants, due to removal of the requirement that to be classified as re-entrants respondents must have worked previously for at least two weeks in a full-time job.

I used three techniques to ensure that my results are not affected by the 1994 changes to the CPS survey. First, in each regression that uses both pre- and post-1993 data, I included a post-1993 dummy variable, so that the results are conditional on the intercept shift associated with the survey redesign; Perry (1972) used a similar approach to account for the 1967 survey redesign. Second, I imposed a direct adjustment to the post-1993 data, as implied by a comparison of the 1993 actual and parallel surveys. The adjustment is based on the percentage change in the total unemployment count and unemployment counts by reason. I also ran all regressions with the post-1993 period excluded. The results for these regressions were all very similar, which indicates that my conclusions are not affected by the 1994 CPS survey redesign.

Estimates of Expected Completed Duration

The CPS data described in the previous subsection provide information on the average length of existing unemployment spells up to the date of the survey. This "average interrupted duration" measure will not in general correspond to the expected duration of a completed spell for a new entrant to unemployment, particularly under changing labor

2. Other studies of unemployment trends have used retrospective data from the March CPS Annual Demographic Supplement, which contains information on individuals' labor force experience during the entire preceding year (e.g., Murphy and Topel 1987). However, monthly data on spells in progress are available on a more timely basis and provide better variation for investigation of time-series properties such as cyclical sensitivity and secular trends. Moreover, as described by a variety of authors, unemployment data are plagued by response biases that are likely to be more severe in retrospective data (Akerlof and Yellen 1982, Levine 1993).

The retrospective data are most useful for joint analyses of unemployment and labor force nonparticipation, as in Juhn, Murphy, and Topel (1991). I focus on unemployment because the two states are behaviorally distinct (see Flinn and Heckman 1983) and unemployment is of independent interest as a macroeconomic variable.

market conditions. Expected completed spell duration depends on the probabilities of continuing in or exiting unemployment as the spell proceeds. Estimation of expected completed durations proceeds as follows.

If the labor market is in steady state—i.e., entry and continuation rates for unemployment spells are constant over time and over the length of a spell—then the total number of unemployed at a particular time can be expressed as the product of incidence and average duration:

$$U = f(0) \cdot D. \tag{1}$$

In (1), U is the number unemployed, D is average expected duration, and $f(0)$ denotes the number of new entrants to unemployment (incidence) in a particular month, which is assumed constant over time in steady state. Then the steady-state estimate of expected completed duration in months, D, is simply the total number of unemployed U divided by the new entrants to unemployment. Using the number of persons unemployed less than five weeks as a measure of the monthly inflow $f(0)$, I can compute this statistic for total unemployment beginning in 1967 and for unemployment by reason beginning in 1976. This is the simplest estimate of expected completed duration based on monthly household survey data.[3]

In contrast, other authors (Sider 1985, Baker 1992a, and Baker, Corak, and Heisz 1998) estimated expected completed duration based on an approach that does not impose steady-state assumptions, which they argue is of particular importance when the cyclical properties of unemployment duration are of interest. The general nonsteady-state approach to estimating expected completed duration using grouped duration data is a "synthetic cohort approach," developed by Kaitz (1970) and Perry (1972).[4] This approach relies on the estimation of monthly continuation rates—i.e., the probabilities that an unemployment spell will continue from one month to the next. These rates in general will vary over the length of a spell due to individual heterogeneity or underlying duration dependence, and they also will vary from month to month as economic conditions change.[5]

Because Sider (1985) and Baker (1992a) started with individual data, they were able to perform relatively exact calculations of unemployment tallies in duration intervals corresponding essentially to the monthly sampling window. This enables relatively precise estimation of the monthly continuation probabilities needed to form a nonsteady-state estimate of expected completed duration.

The monthly BLS data that I use also provide information on unemployment tallies within duration intervals. However, these intervals do not correspond to the monthly sampling window.[6] Moreover, the BLS intervals do not provide adequate information regarding long spells of unemployment. Although such data have been used in the past to form monthly continuation probabilities (e.g., Kaitz 1970), this approach requires substantial data smoothing and reassignment.

I therefore focus on steady-state estimates of expected completed duration. As discussed in the results section, this estimator closely replicates the cyclical properties of Baker's (1992a) nonsteady-state estimator for a comparable sample period. This may seem surprising, given that the nonsteady-state estimator is specifically designed to account for cyclical variability. However, the lagged information used by the nonsteady-state estimator implies that it is not a measure of the expected duration for an individual entering unemployment in the current month.[7] In contrast, although the steady-state estimator requires only information from the current month, the unemployment level used in its denominator reflects much of the same lagged information on continuation rates as is used in the nonsteady-state estimator. Also, Baker (1992b) noted that estimates of expected duration are more sensitive to data smoothing and allocation rules than to steady-state assumptions.[8]

3. An alternative steady-state measure of unemployment duration is the "experience-weighted" measure discussed by Akerlof and Main (1981). It measures the length of the spell to which the average week of unemployment belongs, or the expected duration of in-progress spells. This measure is substantially larger in general than expected completed duration for an individual entering unemployment. Previous authors have not examined the time-series properties of this measure.

4. This is a "synthetic cohort" approach in that with a rotating monthly sample such as the CPS, the estimate of unemployment continuation probabilities is formed by comparing different groups over time, rather than by following the same individuals through time.

5. An alternative approach, which allows for changing continuation or escape rates over the length of a spell (duration dependence), is based on estimating the parameters of a specified distribution of escape rates using data on ongoing spells. Salant (1977) and others used a gamma density to characterize escape rates in such a model. In terms of its approach to changing continuation rates, this model lies between the pure steady-state approach described above and the nonsteady-state approach used by Sider (1985) and Baker (1992a).

6. The published BLS duration intervals are < 5 weeks, 5–10 weeks, 11–14 weeks, 15–26 weeks, 27–51 weeks, and > 51 weeks.

7. Corak and Heisz (1996) propose and estimate a forward-looking nonsteady-state estimator, which reflects the evolution of continuation probabilities into the future for individuals entering unemployment in the current month. They find that their estimator has desirable properties relative to the standard backward-looking nonsteady-state estimator.

8. Although Sider (1985) found substantial differences in results based on nonsteady-state and steady-state estimators, he used the gamma density approach of Salant (1977) rather than the steady-state estimator that I use.

A final estimation issue is "digit preference"—the tendency for measured durations to bunch at week values corresponding to integer multiples of one month and half-years (i.e., multiples of 4 or 26). Previously, analysts handled this problem by allocating a fixed percentage of bunched observations to the next monthly interval. For example, Sider (1985) and Baker (1992a) assigned 50 percent of the bunched observations to the next monthly interval. Baker (1992b) reports that although estimates of expected completed duration are very sensitive to the allocation rule, cyclical elasticity regression results are not. My estimator uses information only on the first monthly interval, and the weekly distribution within this interval is not identified. Given these considerations, I used a modified 50 percent allocation rule based on the grouped interval data. I assume a uniform distribution by weeks within the first monthly interval; the resulting allocation of 50 percent of the implied number at 4 weeks reduces the size of the entrant group by 12.5 percent. This makes my estimates of expected completed duration comparable to Sider's and Baker's; the regression estimates of the cyclical elasticity and time trend are unaffected.

III. RESULTS

Tabulations

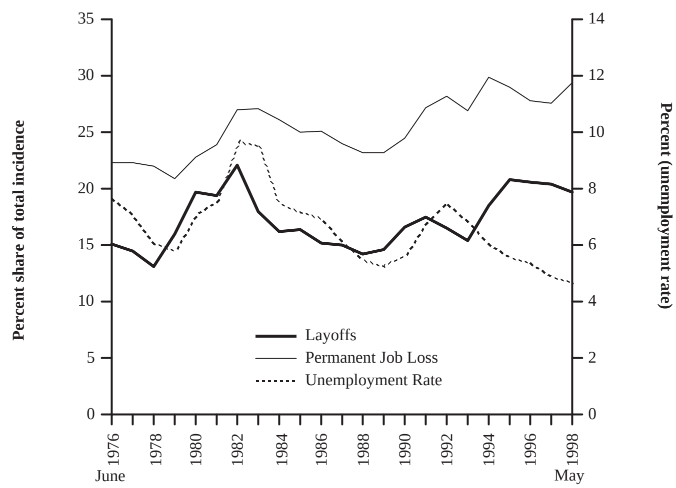
Figures 1–4 show yearly average values of the unemployment rate, unemployment incidence by reason, and unemployment duration (total and by reason);[9] the unemployment rate is identified by the right-hand scale in each figure. Incidence is measured by the number of interrupted spells of less than five weeks duration during the sample month. The values for 1994–1998 in Figures 1–4 reflect adjustments intended to neutralize the impact of the 1994 survey redesign, as described in Section II.

Figure 1 shows the unemployment rate and the shares of layoffs and permanent job loss in total unemployment incidence, each expressed in percentage points, for the period 1976–1998. Job losers on average account for about 43 percent of the newly unemployed during the period, with permanent job losses substantially outnumbering layoffs. Both series appear to be countercyclical, rising and falling with the unemployment rate. However, following the cyclical increase in the early 1990s, permanent job loss has remained very high throughout the decade, which is indicative of an upward trend. Moreover, layoff incidence increased sharply in 1994 and 1995 and has remained high. Overall, the rate of job loss in 1998 is above its sample period average, despite the low unemployment rate prevailing in 1998.

Figure 2 displays unemployment incidence shares for voluntary job leavers (quits) and labor force entrants. Job

FIGURE 1

INCIDENCE OF INVOLUNTARY UNEMPLOYMENT (ANNUAL AVERAGES)



Note: Post-1993 figures corrected for 1994 survey change.

FIGURE 2

INCIDENCE OF VOLUNTARY UNEMPLOYMENT AND LABOR FORCE ENTRY (ANNUAL AVERAGES)

Note: Post-1993 figures corrected for 1994 survey change.

9. The yearly average for 1998 is based on data for the first five months. Also, the incidence by reason figures first became available in June 1976, so the 1976 average is based on the last seven months of the year.

leaving constitutes a relatively small share of unemployment incidence—14 percent on average—but it exhibits a pronounced procyclical pattern. Re-entrant unemployment is frequent and appears to demonstrate moderate procyclicality, although its level has remained low in recent years. New entrant unemployment incidence exhibits limited cyclicality but an apparent downward trend.[10]

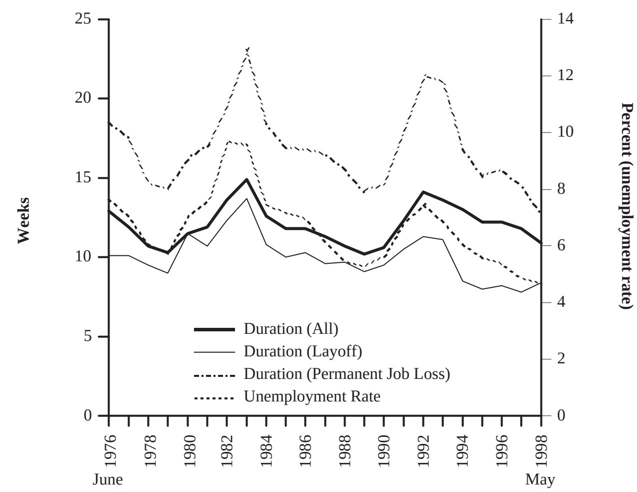
Figure 3 displays yearly average values for the steady-state estimate of expected completed duration of unemployment for all unemployed and by job loss category.[11] Each of these expected duration series exhibits noticeable countercyclicality. Permanent job losers on average endure long spells of unemployment; during the period 1976–1998, the expected duration of unemployment was 17 weeks for permanent job losers and 12 weeks for all unemployed.

Figure 4 shows expected completed duration for all unemployed, quits, and labor force entrants.[12] Individuals unemployed for each of these reasons experience durations around 10 weeks on average, slightly below the overall average, and all appear to exhibit countercyclicality.

Overall, the patterns in Figures 1–4 and underlying tabulations suggest that unemployment spells can be placed in long and short groups according to the reason for unemployment. Workers unemployed due to permanent job loss tend to suffer lengthy spells, with a high degree of countercyclicality in duration. Workers unemployed due to layoffs, voluntary mobility, or labor force entrance typically encounter relatively short spells. An apparent upward trend in unemployment duration may be linked to changes in incidence and duration of unemployment by reason, particularly the rise in permanent job loss. Implementation of more direct tests requires the regression approach described in the next subsection.

FIGURE 3

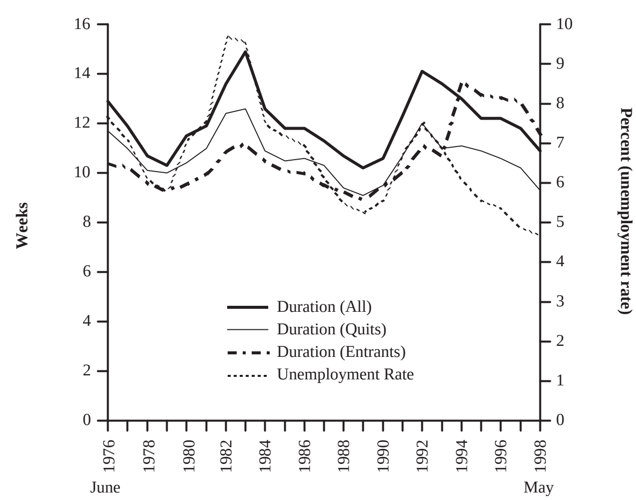
Expected Duration of Unemployment, Total Unemployment, and Job Losers (annual averages)



Note: Post-1993 figures corrected for 1994 survey change.

FIGURE 4

Expected Duration of Unemployment, Total Unemployment, Quits, and Labor Force Entrants (annual averages)



Note: Post-1993 figures corrected for 1994 survey change.

10. In Figure 2, the re-entrant series moves downward sharply between 1993 and 1994 (as does the duration of re-entrant unemployment in Figure 4). These movements may reflect the 1994 survey redesign and my data adjustment intended to overcome it. However, as noted in the text, the regression results reported in the next subsection are robust with respect to the survey redesign.

11. I multiplied estimates of expected duration in months by 4.3 to obtain expected duration in weeks.

12. I merged the re-entrant and new entrant duration series together (weighted by relative incidence), because they are nearly identical over my sample frame (including an upward jump between 1993 and 1994).

Regression Results

Following Perry (1972), Sider (1985), and Baker (1992a), I estimate time-series regression equations to explain the variation in monthly measures of unemployment duration and incidence. My basic approach is to regress these measures on the current unemployment rate (not seasonally adjusted), season dummies, a linear time trend, and a dummy variable that accounts for the CPS survey changes effective after 1993; unemployment duration and the unemployment rate are measured in natural logs.[13] I also report autocorrelation tests for all specifications; the tests are based on Durbin-Watson statistics for models without lagged dependent variables and Durbin's modified test (Durbin 1970) for models that include a lagged dependent variable.

Table 1 presents regression results for the steady-state unemployment duration models. Panel A lists results for total unemployment. The first row lists results for the full sample period of 1967–1998. The results indicate an elasticity of expected duration with respect to the unemployment rate of 0.45 and a small but statistically significant upward time trend. The Durbin-Watson statistic indicates that positive autocorrelation is present in this regression.

In the second row, I included a single period lag of the dependent variable in the regression. This specification is a partial adjustment model, in which the coefficients on the independent variables represent partial or short-run effects; the full long-run effects are obtained by adjusting the coefficients by a factor that is proportional to the coefficient on the lagged dependent variable.[14] Use of this specification in row 2 eliminates autocorrelation as a potential problem. Moreover, the adjusted coefficients indicate a full business cycle elasticity of 0.443 and full time trend effect of 0.026, which are identical to the estimates from the first row.[15] This time trend effect implies an increase in expected duration of about 10 percent between 1967 and 1998.

In rows 3 and 4 of Panel A, I present results for periods beginning in 1976. Both the estimated business cycle elasticity and the time trend are larger for this sample period than for the full sample. The transformed coefficients reveal a long-run business cycle elasticity of 0.57 and a full time trend effect of 0.064. This time trend effect implies nearly a 17 percent increase in the expected duration of unemployment between 1976 and 1998. A separate regression (not reported) verified the absence of a time trend for the period 1967–1980, which explains the substantially larger time trend estimated for 1976–1998 than for the full sample period. In row 4, I estimate the model on data for the years 1976–1993, in order to ensure that the results for the 1976–1998 period are not influenced by inadequate adjustment for the 1994 survey redesign. The results for this restricted sample period are virtually identical to those for the full period from row 3.

The estimated business cycle elasticities in rows 1–4 of Panel A are smaller than Baker's (1992a) estimate of 0.62 for the period 1980–1988, although they are close for the samples that begin in 1976. It is unclear whether this is due to differences across subperiods in the data or to different properties of my steady-state duration estimator and Baker's nonsteady-state measure. Row 5 presents results using my duration estimator and Baker's sample period. The full long-run business cycle elasticity implied by the estimated coefficient is 0.58, which is very close to Baker's estimate. Although full validation of the steady-state approach would require direct comparison of the two approaches for various sample periods, it appears that the steady-state estimator captures cyclical variation reasonably well. Row 5 of Panel B also indicates an upward trend in expected duration during the 1980s, although the shorter sample period reduces the statistical significance of the trend estimate compared to the preceding regressions.

Panel B of Table 1 presents results for expected unemployment duration by reason. The results indicate significant countercyclicality in expected unemployment durations for all reasons, as found by Baker (1992a). The largest cyclical effect is associated with job loss: the transformed coefficients on the unemployment rate in rows 1–3 indicate a business cycle elasticity of about 0.7 in each case. Smaller cyclical effects are evident for labor market entrants; the transformed coefficients imply an elasticity around 0.3.

The time trend effects by reason exhibit less uniformity than the cyclical effects by reason. A substantial upward time trend is evident for unemployment durations experienced by permanent job losers and labor market re-entrants; these time trends imply secular increases of approximately 24 percent and 11 percent in the expected unemployment durations of permanent job losers and labor market re-entrants,

13. This essentially replicates Baker's (1992a) specification. Sider (1985) used deviations from the trend in the index of industrial production (IIP) as his cyclical measure, and he included month dummies in the regression. My results are virtually identical when I replace my season dummies with month dummies. However, I do not report results using the IIP, because my preliminary regressions indicated that its relationship with labor market conditions has changed over time.

14. Specifically, if β is the coefficient on an independent variable and λ is the coefficient on the lagged dependent variable, the full effect of changes in the independent variable is $\beta/(1 - \lambda)$.

15. I estimated the sampling distributions of the transformed coefficients using the standard normal bootstrap approach (Efron and Tibshirani 1993). The resulting estimates of statistical significance are similar to those implied by the unadjusted coefficients and standard errors, so I do not list the sampling statistics for the transformed coefficients.

TABLE 1

Unemployment Duration Regressions

Dependent Variable	ln (unemployment rate)	Time Trend (x 100)	Lagged Dependent Variable	Autocorrelation Test[a]
		Panel A: Total Unemployment		
ln(duration), 1967–98	0.446**	0.026**	—	DW = 1.64
	(0.021)	(0.006)	—	
ln(duration), 1967–98	0.330**	0.019**	0.255**	–0.069
	(0.033)	(0.006)	(0.057)	(0.105)
ln(duration), 1976–98	0.450**	0.051**	0.206**	–0.255*
	(0.053)	(0.011)	(0.073)	(0.126)
ln(duration), 1976–93	0.429**	0.050**	0.259**	–0.082
	(0.055)	(0.010)	(0.078)	(0.137)
ln(duration), 1980–88	0.474**	0.070*	0.189	0.041
	(0.083)	(0.032)	(0.112)	(0.196)
		Panel B: Unemployment by Reason, 1976–1998		
ln(duration), total job losers	0.382**	0.050**	0.462**	–0.244*
	(0.049)	(0.011)	(0.057)	(0.091)
ln(duration), temporary layoffs	0.376**	0.022	0.469**	0.166
	(0.058)	(0.012)	(0.068)	(0.106)
ln(duration), permanent job losers	0.418**	0.055**	0.408**	–0.518**
	(0.053)	(0.012)	(0.057)	(0.093)
ln(duration), job leavers	0.339**	0.003	0.227**	–0.380**
	(0.049)	(0.011)	(0.068)	(0.141)
ln(duration), re-entrants	0.284**	0.041**	0.021	–0.478*
	(0.045)	(0.012)	(0.073)	(0.204)
ln(duration), new entrants	0.251**	–0.014	0.187*	–0.466*
	(0.064)	(0.018)	(0.075)	(0.198)

Note: Standard errors are in parentheses. Other explanatory variables included in all regressions are three season dummies and a CPS survey redesign dummy for regressions that include years beyond 1993 (post-1993 = 1). The number of observations is 376 for 1967–98, 263 for 1976–98, 210 for 1976–93, and 108 for 1980–88. The 1976 data begin in June.

* Significant at the 5% level, two-tailed test.

** Significant at the 1% level, two-tailed test.

[a] The test statistics are the Durbin-Watson statistic or the t test on the listed regression coefficients for single-period lagged residuals.

respectively, over the period 1976–1998. The expected duration of temporary layoff unemployment also appears to have risen somewhat; the transformed coefficient is significant at the 5 percent level and implies an increase in expected duration of 11 percent between 1976 and 1998.[16]

Table 2 reports results for linear probability models of unemployment incidence by reason.[17] Layoffs and other job losses are strongly countercyclical; job leaving and labor market re-entrance is procyclical; and new labor

16. The presence of negative autocorrelation in these models implies that the standard error estimates are conservative (i.e., the associated t tests probably understate significance levels).

17. Baker (1992a) reports no difference between results based on linear probability and logistic models in this setting, probably because the incidence variables all are sufficiently well-bounded away from zero and one.

TABLE 2

REGRESSIONS OF UNEMPLOYMENT INCIDENCE (SHARES) BY REASON, 1976–1998

DEPENDENT VARIABLE	ln (unemployment rate)	TIME TREND (x 100)	LAGGED DEPENDENT VARIABLE	AUTOCORRELATION TEST[a]
% TOTAL JOB LOSERS	0.108**	0.019**	0.265**	0.137
	(0.021)	(0.005)	(0.080)	(0.116)
% temporary layoffs	0.041**	0.000	0.356**	0.222*
	(0.013)	(0.003)	(0.066)	(0.097)
% permanent job losers	0.069**	0.022**	0.014	–0.402*
	(0.010)	(0.003)	(0.073)	(0.162)
% JOB LEAVERS	–0.062**	0.000	0.454**	0.397**
	(0.011)	(0.002)	(0.073)	(0.125)
% RE-ENTRANTS	–0.037**	–0.014**	0.039	0.303*
	(0.010)	(0.003)	(0.068)	(0.153)
% NEW ENTRANTS	0.003	–0.010**	0.034	0.402**
	(0.010)	(0.003)	(0.075)	(0.104)

NOTE: Standard errors are in parentheses. Other explanatory variables included in all regressions are three season dummies and a CPS survey redesign dummy (post-1993 = 1). The number of observations is 263.
* Significant at the 5% level, two-tailed test.
** Significant at the 1% level, two-tailed test.
a The test statistics are the t tests on the listed regression coefficients for single-period lagged residuals.

market entrants do not respond to business cycle conditions. The transformed coefficients imply long-run elasticities with respect to the unemployment rate that are broadly comparable to those obtained by Baker (1992a), although my estimates are somewhat smaller in general. I also find a significant upward time trend in the incidence of permanent job loss, consistent with results using individual panel data in Valletta (1998). The transformed trend coefficient indicates nearly a 6 percentage point increase in the incidence of permanent job loss between 1976 and 1998; this equals nearly a 25% increase relative to the sample mean incidence of permanent job loss. The results also reveal downward trends in the incidence of labor market entrant unemployment; the coefficients indicate approximately 10 percent and 20 percent declines between 1976 and 1998 in the incidence of re-entrant and new entrant unemployment, respectively.[18]

Decomposition Analysis

Three key results arising from the analysis thus far are: (1) longer durations of unemployment for permanent job losers than for other groups; (2) secular increases in total unemployment duration and duration for job losers; (3) rising incidence of permanent job loss to unemployment. It seems likely that rising duration of total unemployment in (2) is due in large part to the change in the composition of unemployment implied by (1) and (3).

I perform several decomposition analyses to investigate this link. I first apply a simplified variant of the decomposition that Baker (1992a) used to test the heterogeneity hypothesis of cyclical variability in unemployment durations, using unemployment by reason as my measure of heterogeneity.[19] Recall from equation (1) that expected unemployment duration in a particular month is the number unemployed divided by the number of new entrants to unemployment during the month. Then expected total unemployment duration equals a weighted average of expected

18. Positive autocorrelation is evident for all but the permanent job loser and total job loser series. However, estimation using the Newey and West (1987) approach to account for autocorrelation indicated that the impact on the estimated standard errors is minimal.

19. Baker also performed decompositions by region, industry, education, and demographic groups. These other decompositions provided even less evidence in favor of the heterogeneity hypothesis than did the decomposition by reason for unemployment.

duration by reason, with the weights equal to the shares of unemployment incidence by reason. This property enables decomposition of total unemployment duration into two components:

D_{pc} ("probability constant")—expected total duration holding expected duration for each reason at its sample average, but allowing the shares of unemployment incidence by reason to change

D_{sc} ("share constant")—expected total duration holding the shares of unemployment incidence by reason equal to their sample averages, but allowing expected duration by reason to change.

Comparison of regressions using the constructed variables D_{pc} and D_{sc} with regressions using the unadjusted duration measure indicates the relative roles of changing duration by reason and changing shares by reason in the determination of the time-series properties of total unemployment duration. These results are listed in Panel A of Table 3. The first row repeats the results for the unadjusted duration measure (row 3 from Table 1, Panel A). The second row lists the results for the probability constant measure D_{pc}, and the third row lists the results for the share constant measure D_{sc}. A comparison of the results in the final two rows indicates that virtually all of the cyclical variability in total unemployment duration is due to cyclical variability in expected duration by reason rather than variability in incidence by reason: the coefficient on the unemployment rate is very small for D_{pc}, which holds expected duration by reason constant, and large for D_{sc}, which holds incidence by reason constant. Most of the upward time trend also is attributable to rising duration by reason. However, about 20 percent of the upward trend in total duration is due to changing incidence by reason.

The decomposition listed in Panel A of Table 3 groups all reasons for unemployment together. Recall, however, that the key changes over time have been in the incidence and duration of unemployment associated with permanent job loss. The first row of Panel B lists results from an alternative decomposition that focuses on permanent job loss. I formed the dependent variable used in the first row of Panel B by holding the incidence and duration of permanent job loss constant at their respective sample averages. Comparison of these results with the results in the first row of Panel A reveals the effect on total expected du-

TABLE 3

Unemployment Duration Regressions, 1976–1998, Adjusted by Reason for Unemployment

Dependent Variable	ln (unemployment rate)	Time Trend (x 100)	Lagged Dependent Variable	Autocorrelation Test[a]
Panel A: Decomposition, Probability and Share Constant, All Reasons				
ln(D) (unadjusted)	0.450**	0.051**	0.206**	–0.255*
	(0.053)	(0.011)	(0.073)	(0.126)
ln(D_{pc}) (probability constant)	0.034**	0.013**	0.151*	–0.517**
	(0.005)	(0.002)	(0.069)	(0.166)
ln(D_{sc}) (share constant)	0.388**	0.037**	0.247**	–0.302*
	(0.047)	(0.010)	(0.070)	(0.122)
Panel B: Incidence and Duration by Reason Held Constant				
ln(D), perm. job loss constant	0.176**	–0.006	0.112	–0.086
	(0.023)	(0.005)	(0.075)	(0.155)
ln(D), entrant constant	0.384**	0.061**	0.184*	0.062
	(0.047)	(0.010)	(0.080)	(0.119)

Note: Standard errors are in parentheses. Other explanatory variables included in all regressions are three season dummies and a CPS survey redesign dummy (post-1993 = 1). The number of observations is 263.

* Significant at the 5% level, two-tailed test.

** Significant at the 1% level, two-tailed test.

[a] The test statistics are the t tests on the listed regression coefficients for single-period lagged residuals.

ration of rising incidence and duration of unemployment due to permanent job loss. The substantially smaller coefficients in row 1 of Panel B than in row 1 of Panel A indicates that rising duration and incidence of unemployment due to permanent job loss accounts for most of the cyclical effect and all of the time trend effect on total duration. In conjunction with the decomposition results from Panel A, which showed that changing incidence explains only a small portion of the time trend, the Panel B results indicate that rising duration associated with permanent job loss has played the dominant role in the trend toward rising duration of total unemployment.

Finally, recall that Tables 1 and 2 also revealed significant changes in unemployment duration and incidence for re-entrants and new entrants to the labor force. The dependent variable in the final row of Panel B in Table 3 holds the incidence and duration of entrant unemployment constant. Comparison of these results with the total duration results from row 1 of Panel A reveals that labor force entrants account for only a small portion of the cyclical variability in total expected duration. Moreover, trends in the incidence and duration of labor force entrant unemployment reduced rather than increased total expected duration over time: the estimated time trend effect is larger when entrant unemployment duration and incidence are held constant (Panel B, row 2) than when they are allowed to vary (Panel A, row 1). This result occurs because the upward trend in expected duration for re-entrants (Table 1, Panel B) is offset by declining incidence for re-entrants (Table 2).

IV. Conclusions

Using steady-state measures of expected duration of unemployment, I find strong countercyclicality in unemployment durations during the period 1967–1998, which is relatively uniform across the different reasons for entry into unemployment. This result updates the previous literature into the 1990s. However, I also uncover an upward trend in duration that raised expected durations by approximately 17 percent between the years 1976 and 1998 (conditional on the unemployment rate as a measure of business cycle conditions). Like previous researchers, I also found substantial cyclical variability in unemployment incidence by reason for unemployment. I extend previous research by uncovering an upward trend in the incidence of permanent job loss and downward trends in the incidence of labor force entrant employment between 1976 and 1998. A decomposition of the increase in overall duration reveals that rising incidence and duration of unemployment due to permanent job loss can account for the full time trend effect during this period, with rising duration playing the dominant role. The incidence and duration of unemployment for other reasons made essentially no contribution to rising total unemployment duration.

Given the potential welfare losses associated with lengthy unemployment spells, the secular trend toward rising unemployment duration merits additional investigation. A useful research goal might be to identify the underlying economic forces that generate the link between rising permanent job loss and rising unemployment durations. In very recent work, Baumol and Wolff (1998) argue that technological progress can lengthen unemployment duration by increasing employment churning and skill mismatches. Their estimates suggest that several measures of technological change in the workplace are strongly associated with increases in the average interrupted spell duration between 1971 and 1994. However, drawing a reliable inference of a causal link between these two phenomena requires additional research. Other possible explanations of the trend toward rising duration include changing job search strategies by job losers and measurement issues related to movements between labor force states. Such hypotheses should be tested directly.

Depending on the underlying cause, the trend toward rising unemployment duration may have macroeconomic implications. Perry (1970) suggested that for a given unemployment rate, differences in the demographic structure of unemployment imply different degrees of wage pressure, since the value of output on the job varies across groups. In his comments on Perry's paper, Solow (1970) argued instead that the duration structure of unemployment may be more important than the demographic structure: "People who have been unemployed a long time put more downward pressure on wages because they are more willing to undercut going wage rates in order to get a job" (p. 445). The model of Blanchard and Diamond (1994) formalizes this reasoning in a matching model of the labor market, and Duca (1996) presents evidence suggesting that longer unemployment durations help explain low wage and price inflation in the early 1990s. This interpretation of recent weak wage pressure is reinforced by the key role of permanent job loss in explaining lengthening durations: research on reemployment prospects of displaced workers indicates that they suffer larger wage losses upon reemployment than do workers who changed jobs for other reasons. Thus, the secular trend toward rising permanent job loss and rising durations may help to explain limited upward wage and price pressure in the current lengthy expansion.

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