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ISSN 0164-0682

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Jeffrey R. Campbell and Zvi Hercowitz

This article uses a calibrated general-equilibrium model of lending from the wealthy to the middle class to evaluate the effects of tightening household lending standards. The authors simulate a rise in down payment and amortization rates from their average values in the late 1990s and early 2000s to levels more typical of the era before the financial deregulation of the early 1980s. Their results show a drop in loan demand. This substantially lowers interest rates for an extended period. Counterintuitively, tightening lending standards makes borrowers better off.

14 Measuring the equilibrium real interest rate

Alejandro Justiniano and Giorgio E. Primiceri

The equilibrium real interest rate represents the real rate of return required to keep the economy's output equal to potential output. This article discusses how to measure the equilibrium real interest rate, using an empirical structural model of the economy.

Second Quarter

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Daniel Aaronson, Bhashkar Mazumder, and Shani Schechter

This article analyzes what is behind the recent unprecedented rise in long-term unemployment and explains what this rise might imply for the economy going forward. In particular, the authors attribute the sharp increase in unemployment duration in 2009 to especially weak labor demand and, to a lesser degree, extensions in unemployment insurance benefits.

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Interest rates following financial re-regulation

Jeffrey R. Campbell and Zvi Hercowitz

Introduction and summary

Mortgages and other forms of household borrowing typically require collateral, such as a house or car. Typical loan contracts require borrowers to take an initial equity stake in the collateral (the down payment) and to increase ownership further by repaying the loan's principal before the collateral fully depreciates (amortization). Since the New Deal, government regulation has substantially influenced these terms of private contracts. In the 1940s and early 1950s, the Federal Reserve Board imposed stringent minimum down payment rates and maximum amortization periods for home mortgages, auto loans, and loans to purchase other consumer durable goods. The suspension of these regulations in 1953 allowed consumer credit to grow steadily until the credit crunch of August 1966. The financial deregulation wave of the late 1970s and early 1980s triggered innovations in consumer lending that further decreased households' ownership stakes in their housing and other tangible property. Many observers have blamed precisely this deregulation for the most recent financial crisis, so it seems very possible that households' required ownership stakes will be rising as policymakers look at their options for improving the regulation of consumer loans and other financial contracts.

In this article, we employ a model of lending from the wealthy to the middle class to evaluate the effects of raising the equity requirements of household debt. We build on our earlier analysis of the Carter–Reagan financial deregulation in Campbell and Hercowitz (2009). In that article, we found that lowering equity requirements raises the demand for household credit and thereby increases the interest rate. This resembles the simultaneous increases in household debt and interest rates during the mid-1980s, even though we abstract from rising government deficits, which are the standard explanation for that period's high interest rates. In this

article, we examine the implications of reversing this process by increasing down payment rates for new loans and by forcing all loans to amortize faster. The model's results show that this reform reduces loan demand. The interest rate falls 78 basis points over three years and then very slowly returns to its level before the reform. In an alternative version of our model in which producers cannot absorb the capital freed by tightening household lending standards, the interest rate falls 129 basis points over the three years after the reform. These results are potentially of interest to monetary policymakers because they can guide an assessment of how financial market reforms impact the “neutral” interest rate required to keep the economy's output at its potential in the absence of business cycles.

In the model, saving households are rentiers living off of their wealth, so the low interest rate unambiguously harms them. Nevertheless, the low rate has two beneficial effects for borrowers. First, the lower interest rate reduces the carrying cost of debt. Second, the lower interest rate brings down the user cost of capital and thereby encourages investment. These investments increase the demand for labor and thereby raise wages. Overall, the model's predictions show that borrowers' welfare gains are equivalent to raising their consumption permanently by 0.9 percent. If we treated the household credit market in isolation from the rest of the economy, then this second effect would be absent.

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In fact, such a market-by-market analysis would be misleading; the reform makes borrowers slightly worse off after shutting down its indirect effect on wages.

If tighter lending standards changed neither the interest rate nor wages, then they must harm borrowers by limiting their choices. Following this intuition about the “direct” effects alone leads to the conclusion that tighter lending standards primarily harm borrowers. Our results show that this intuition can easily be overturned by a complete equilibrium analysis that accounts for the “indirect” effects of changing prices. Since the reform helps some households at the expense of others, its assessment requires us to weight the households’ utility changes. Even with a specific assumption about these weights, the result is only a partial assessment, since we have nothing to say about how tightening household lending standards changes systemic economic risk.

Our article proceeds as follows. In the next section, we review the history of interest rates and household debt markets in the United States, paying particular attention to households’ ownership stakes in their tangible property. Then, we present the model and derive its long-run implications for debt and interest rates. We show that financial re-regulation has no long-run effect on interest rates, leaves saving households worse off, and improves borrowers’ welfare. Finally, we present the complete analysis of the reform.

Household debt and interest rates in the United States

The rise of mass production techniques early in the twentieth century created a large volume of standardized capital goods, such as automobiles, which could serve as collateral for credit extended to households. By the 1920s, most durable household goods could be bought “on credit” directly from their retailers. The home mortgage market of that decade bears a remarkable resemblance to that of the 1990s and 2000s. First mortgages had low loan-to-value ratios, and households often financed the first mortgage’s required down payment with second and third mortgages. All of these mortgages matured in only a few years, and they required no repayment of principal before maturing.¹

The Great Depression, World War II, and the Korean War dramatically increased government involvement in consumer credit markets. In the early 1930s, the federal government purchased large volumes of “underwater” mortgages. These were loans with principals exceeding the value of their collateral. It then refinanced them with 15-year amortized mortgages, which built in the gradual repayment of the principal over the 15-year amortization period. This amortization directly served the policy goal of avoiding a wave of mortgage

defaults arising from a sudden lack of refinancing options. The 15-year amortized mortgage and its 30-year cousin accounted for most household debt from the 1930s through the 1980s, even though they required substantial down payments from borrowers.² The move from interest-only short-term loans to long-term amortized debt reduced systemic risk at the cost of keeping potential homeowners with insufficient funds for a mortgage’s down payment out of the market. With the onset of World War II, the Federal Reserve Board tightened loan standards further by issuing Regulations X and W. These dictated restrictive maximum loan-to-value ratios and amortization periods for home mortgages (Regulation X) and other collateralized consumer credit (Regulation W).

The Federal Reserve suspended enforcement of Regulations X and W near the end of the Korean War in 1953. Figure 1 illustrates the evolution of credit markets since 1952. The data come from the Federal Reserve Board’s *Flow of Funds Accounts of the United States*. The dashed line in figure 1 shows the ratio of all mortgage debt on owner-occupied housing relative to this housing stock’s value, and the solid line represents the ratio of all household debt to all tangible assets of households, which include the stock of owner-occupied real estate and the stock of automobiles owned by households. Since these are both useful measures of household leverage (the use of debt to finance investment), we refer to them henceforth as *leverage ratios*.

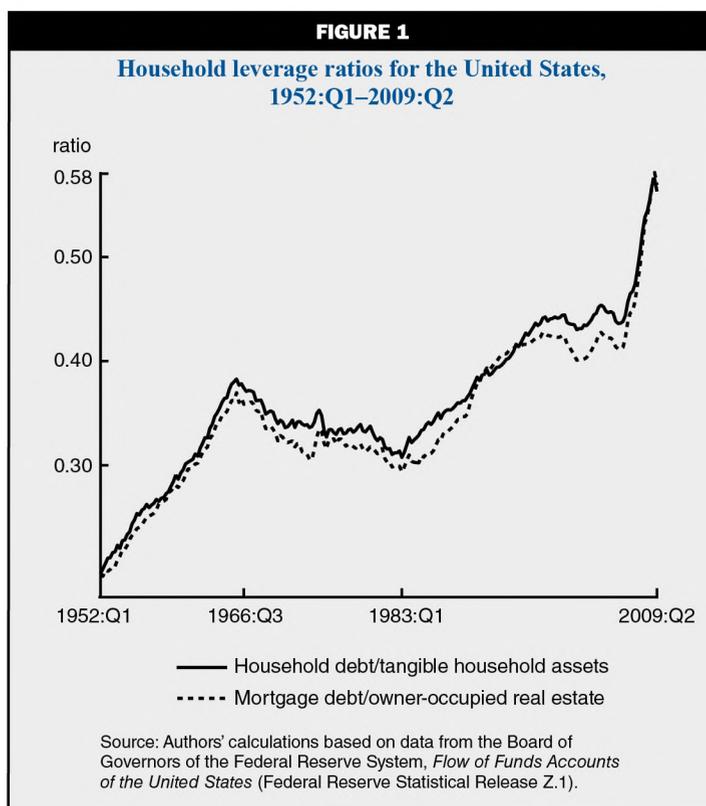
The wartime credit restrictions made these leverage ratios very low: They both equal about 0.195 in the first quarter of 1952. Throughout the 1950s, both ratios rise dramatically. The overall leverage ratio (the solid line in figure 1) peaks at 0.38 in the fourth quarter of 1965. At that time, the Federal Reserve’s Regulation Q placed a cap on the permissible interest rate paid on savings accounts. During the credit crunch of August 1966, market interest rates exceeded this cap, and the resulting outflow of funds from savings and loans and other traditional sources of mortgages reduced the availability of household credit.

The mid-1960s marked a turning point for household leverage ratios. They declined (not always steadily) until the enactment of the Garn–St Germain Depository Institutions Act in the last quarter of 1982. This act and the Monetary Control Act of 1980 eliminated many restrictions on mortgage lending. Along with the concurrent growth of mortgage debt securitization, these changes fueled a second wave of post-war household leverage growth. In the first quarter of 1983, both ratios equaled about 0.30. By the first quarter of 1995, they both equaled 0.41.

Throughout the credit expansion of the late 1990s and the early 2000s, these ratios rarely exceeded 0.45. Home prices began to decline in the middle of 2006, mechanically raising the household leverage ratios. This continued until the first quarter of 2009, when both ratios equaled about 0.58. The most recently available data come from the second quarter of 2009, and they show the leverage ratios declining. Of course, the leverage ratios' common recent spike emanated from a loss in the value of previously mortgaged properties rather than from any deliberate loosening of mortgage terms. With their mortgages considered underwater, many homeowners chose to delay repayment or default outright. The financial turmoil that arose from the resulting impairment of mortgage debt has led most observers to reassess the need for tighter mortgage standards. Therefore, we expect these household leverage ratios to continue their declines as creditors write off their bad debts (thus reducing household indebtedness) and as lenders raise required down payments and principal repayment rates on newly issued loans. Furthermore, the possibility of congressionally mandated changes to financial market regulation might either directly or indirectly lead to tighter standards for household credit.

We expect tighter loan standards to reduce demand for credit, thereby lowering interest rates. To get a sense of how much lower we could expect them to go, we plot the yield on three-year constant-maturity zero-coupon U.S. Treasury debt in figure 2. To account for the effects of anticipated inflation on these interest rates, we have subtracted from each of them the most recent four-quarter percentage change in the Personal Consumption Expenditures Price Index. The yield's average over the time period plotted (the fourth quarter of 1953 through the third quarter of 2009) is 2.6 percent.

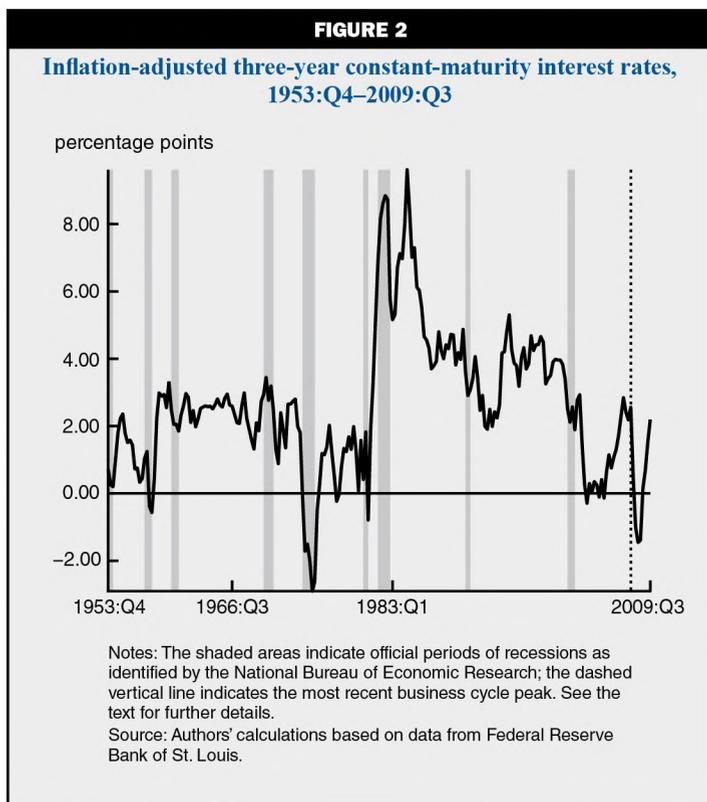
The most noticeable feature of the data is the familiar rise in real interest rates associated with the Federal Reserve's policy of targeting the growth rate of money that began in the fourth quarter of 1979 and ended in the fourth quarter of 1982. To get a better sense of the relationship between credit demand and interest rates, we remove this period and that of the recent financial crisis from the analysis. For the remainder, we have calculated average interest rates for the periods defined by turning points of the household leverage ratios in figure 1: These are 1953:Q4–1966:Q3,



1966:Q4–1979:Q3, 1983:Q1–1995:Q4, and 1996:Q1–2007:Q2. The results are 1.94 percent, 1.33 percent, 4.50 percent, and 2.45 percent. Thus, it appears that the interest rate rose at the same time household leverage ratios were growing in the 1980s and early 1990s, and a decline in interest rates accompanied the end of both growth spurts in figure 1. An explanation of interest rates that focuses only on household leverage ratios is clearly incomplete. For example, contemporaries attributed the high interest rates of the 1980s to that era's high government deficits.³ Nevertheless, the association between interest rates and changes in household leverage seems strong enough to merit further quantitative exploration. We next present a theoretical framework for doing so.

A model of household debt and interest rates

Much of modern macroeconomic theory builds on the useful fiction that identical infinitely lived households populate the economy. This will not do for the question at hand because two identical households have no incentive to lend to each other. Accordingly, our model of household debt and interest rates has two representative households, which we call the *borrower* and the *saver*. The borrower is less patient



than the saver. The difference in patience motivates the (heads of) households to live up to the names we have assigned them. If the borrower's debts were limited only by her ability to repay them, then she would never stop borrowing more. As time passes, she would spend more and more on interest payments and less and less on her own consumption.⁴ This is grossly unrealistic for the United States as a whole. Another feature of our model—collateral requirements—inhibits the never-ending expansion of debt. In the long run, the saver's consumption-savings decisions determine the interest rate. At that rate, the borrower would like to expand her debts. However, the collateral requirement inhibits her from doing so.

As noted previously, most household debts require the borrower to hold an equity stake in the good serving as collateral. The borrower's down payment is the equity stake at purchase, and the equity stake grows as the borrower repays the loan's principal. In the model, two parameters determine the borrower's equity requirements. Because the history of household debt in the United States indicates that government regulation substantially influences equity requirements, we view the two equity requirement parameters as being set by policy.⁵ In Campbell and Hercowitz (2009), we modeled the expansion of leverage following the financial market deregulation of the early 1980s as a reduction

of government-set equity requirements. To consider the effects of anticipated increases in equity requirements on the interest rate, we now reverse that experiment by raising the equity requirement parameters.

Next, we present the model of household debt and interest rates. We begin by describing the two households' preferences. We then lay out the economy's technology for producing goods, and we finish with a discussion of both households' consumption and savings choices in a competitive equilibrium.

Consumer choices

Both the saver and the borrower value the consumption of two goods. The first good is nondurable and stands in for items such as food, energy, and entertainment services. The second good represents the use of durable goods such as housing, furniture, automobiles, and consumer electronics. Both individuals can adjust their consumption of these goods once every calendar quarter.

We denote the quantity of the nondurable good consumed in quarter t by the borrower with \hat{C}_t . The analogous quantity for the saver is \tilde{C}_t . Similarly, we represent the quantities of the durable good used by the borrower and saver in quarter t with \hat{S}_t and \tilde{S}_t . Henceforth, we use \hat{A} and \tilde{A} to represent borrower- and saver-specific versions of A .

If these households are to make consumption and savings decisions, then they need to know how to trade off nondurable and durable consumption in the present quarter and how to balance consuming more of either good today versus saving to enable more consumption in the future. For this, we suppose that they plan how much of both goods to consume in the present quarter and in every future quarter. We denote a plan for the borrower's nondurable consumption from quarter t onward with $\hat{C}^t = (\hat{C}_t, \hat{C}_{t+1}, \hat{C}_{t+2}, \dots) = (\hat{C}_t, \hat{C}^{t+1})$. The borrower's analogous plan for durable consumption is $\hat{S}^t = (\hat{S}_t, \hat{S}_{t+1}, \hat{S}_{t+2}, \dots) = (\hat{S}_t, \hat{S}^{t+1})$. We suppose that for each possible plan, the borrower computes a utility value $\hat{U}(\hat{C}^t, \hat{S}^t)$, using the following formula:

$$\hat{U}(\hat{C}^t, \hat{S}^t) = \theta \ln \hat{S}_t + (1 - \theta) \ln \hat{C}_t + \beta \hat{U}(\hat{C}^{t+1}, \hat{S}^{t+1}).$$

The parameters θ and β both lie between zero and one. This says that the utility value of following a plan equals the *felicity* from the current quarter's

consumption, $\theta \ln \hat{S}_t + (1 - \theta) \ln \hat{C}_t$, plus the value of continuing to follow the plan discounted by $\hat{\beta}$.⁶

The saver's utility value of a given plan can be calculated from his analogous equation:

$$\tilde{U}(\tilde{C}^t, \tilde{S}^t) = \theta \ln \tilde{S}_t + (1 - \theta) \ln \tilde{C}_t + \tilde{\beta} \tilde{U}(\tilde{C}^{t+1}, \tilde{S}^{t+1}).$$

The value of θ here equals its value in the borrower's utility rule, so both households agree on how to divide an allocation of income for the current quarter between nondurable goods and (the services from) durable goods to make felicity as large as possible. However, the saver's discount factor $\tilde{\beta}$ exceeds the borrower's discount factor $\hat{\beta}$. In this sense, the borrower is less patient than the saver. The borrower would prefer to trade the saver's best possible consumption plan for one of equal cost, but with higher consumption in the present and lower consumption in the future.

Production of income and accumulation of wealth

Each quarter, the economy inherits three stocks of capital goods from the previous quarter. The first is the stock of *market capital*. We denote the number of machines in the stock of market capital available in quarter t with K_t . Combining these machines with N_t hours of work (provided in principle by either household) yields an output of $Y_t = K_t^\alpha N_t^{1-\alpha}$, measured in units of the nondurable consumption good. After production, a fraction λ of the machines stop working. Investments in machines, I_t , can replace those lost to depreciation and (if sufficiently large) expand the stock of machines available for the next quarter. Thus,

$$K_{t+1} = (1 - \lambda)K_t + I_t.$$

The remaining two stocks inherited from the previous quarter are the two households' stocks of *home capital*, that is, durable goods. We assume that the flow of services from a stock of home capital is proportional to its size, so that we use \hat{S}_t and \tilde{S}_t to represent each of the households' durable goods stocks as well as the flows of services forthcoming from them. Just as with market capital, the home capital goods depreciate and can be replaced and expanded with investment. The two stocks' common depreciation rate equals δ , and their respective investments are \hat{X}_t and \tilde{X}_t . Therefore,

$$\hat{S}_{t+1} = (1 - \delta)\hat{S}_t + \hat{X}_t,$$

and

$$\tilde{S}_{t+1} = (1 - \delta)\tilde{S}_t + \tilde{X}_t.$$

All income in the economy can be directed toward one of the following uses: each household's nondurable consumption, investment in each household's stock of home capital, or investment in the stock of market capital. It is costless to convert one unit of income into one unit of any of these goods. Since the uses of income cannot exceed that available, we have

$$\hat{C}_t + \tilde{C}_t + \hat{X}_t + \tilde{X}_t + I_t \leq Y_t.$$

The households face two other substantial limits on their accumulation of capital. First, the machines in the stock of market capital may not be converted into consumption goods of either kind. This makes sense for most capital goods because blast furnaces and airliners are of little use to the typical consumer. We impose this limit by requiring that $I_t \geq 0$. Second, neither household may sell durable goods from their stocks of home capital. That is, $\hat{X}_t \geq 0$ and $\tilde{X}_t \geq 0$. Obviously, households can and do sell their durable goods all of the time. However, we find this assumption reasonable when we suppose that the model's borrower and saver represent two classes of individuals with different tastes. If the saver is rich and consumes mansions while the borrower is middle class and consumes bungalows, then the restriction means that we cannot convert mansions into bungalows and vice versa.

Trade and competition

We have now described how the two households rank consumption plans and the technology available for implementing them. We will now present how the households implement these plans by reviewing a typical quarter's trades in the sequence they occur. We then describe the collateral requirements that restrict the households' debts and finish with a presentation of the conditions required for markets to clear.

The sequence of trades in a quarter

At the beginning of quarter t , the households own their stocks of durable goods; stocks of market capital, \hat{K}_t and \tilde{K}_t ; and financial assets (bonds), \hat{B}_t and \tilde{B}_t .

Production takes place at a representative firm. It rents capital from the households and combines it with labor to produce income. The cost of renting one machine in quarter t is H_t , and the cost of one hour of

work is W_t . Capital and labor employed at each firm are chosen to maximize its profits. After production takes place, the representative firm makes its required rental payments to the owners of capital, returns the undepreciated capital goods to their owners, and pays its wage bill. We think of the saver as representing the wealthiest families in the United States, so we suppose that he spends all of his time on leisure activities and offers none to the labor market. The borrower represents the middle class, so we suppose that she offers N hours of work to the market regardless of the wage she earns for each one. Thus, the saver's wage income equals zero always, while the borrower's is $W_t N$.

The funds available to either household is the sum of that household's labor earnings, the rents it receives for the use of its market capital, and its stock of bonds. It can put these funds to one of four uses. Three of these—nondurable consumption, investment in home capital, and investment in market capital—have already been covered. The fourth use of funds is the purchase of new bonds. All bonds pay one unit of the nondurable consumption good in the next quarter, and their price in the current quarter is $1/R_t$, where R_t is the gross rate of interest. With this in place, we can write the two households' budget constraints as

$$\hat{C}_t + \hat{X}_t + \hat{I}_t + \hat{B}_{t+1} / R_t \leq W_t N + H_t \hat{K}_t + \hat{B}_t$$

and

$$\tilde{C}_t + \tilde{X}_t + \tilde{I}_t + \tilde{B}_{t+1} / R_t \leq H_t \tilde{K}_t + \tilde{B}_t$$

Collateral requirements

A household can choose any positive value of bonds (\hat{B}_{t+1} or \tilde{B}_{t+1}) that is consistent with its budget constraint. When either of these bond stocks is *negative*, we say that household is indebted. An indebted household must pledge some or all of its home capital stock as collateral. We denote the maximum debts that can be collateralized by the two households' home capital stocks with \hat{V}_t and \tilde{V}_t . So, we require:

$$-\hat{B}_t \leq \hat{V}_t$$

and

$$-\tilde{B}_t \leq \tilde{V}_t$$

We specify these maximum debts with

$$\hat{V}_{t+1} = (1 - \phi) \hat{V}_t + (1 - \pi) \hat{X}_t$$

and

$$\tilde{V}_{t+1} = (1 - \phi) \tilde{V}_t + (1 - \pi) \tilde{X}_t$$

Here, $1 - \pi$ is the maximum loan-to-value ratio allowed for household debt, and ϕ is the rate at which the principal must be repaid. If $\phi = \delta$, then the borrower must repay the principal only to the extent that depreciation erodes the collateral's value. If instead $\phi > \delta$, then the borrower must accumulate equity in the collateral as it ages. We adopt the specification requiring the geometric repayment of principal because it greatly simplifies the ensuing analysis.

Market clearing and equilibrium

The evolution of the model economy can be completely described by a collection of plans for current *and future* nondurable consumption, durable consumption, market capital, and collateral values, as well as the sequences of the wage rate, the rental rate of capital, and the interest rate. We say that such a collection is an *equilibrium* if the households' consumption plans maximize their utility values given their incomes; the representative firm maximizes its profit given the wage and interest rate; and the demands for bonds, market capital, and labor always equal their corresponding supplies. The interested reader can find a more technical definition of equilibrium in box 1.

The model's steady state

Next, we examine how the steady-state values of the model's key outcomes change with parameters so that we can gain intuition valuable for interpreting the model's dynamics. By definition, a steady state is an equilibrium in which all of the variables are constant over time. Therefore, a household's borrowing constraint binds either always or never. It is tedious but not difficult to show that only the less patient household's borrowing constraint binds in the steady state.

For our purposes, the three key variables of interest are the interest rate and the two households' leverage ratios (their stocks of outstanding household debts divided by the values of their household capital stocks). To characterize all of these variables, we first need to consider both households' optimal consumption and savings choices. Suppose that the saver begins with a utility-maximizing steady-state consumption plan with nondurable consumption \tilde{C} and changes it slightly by decreasing consumption in year t by $\Delta > 0$,

BOX 1

Equilibrium definition

Building upon the notation we used for the two households' consumption plans, we denote the path for any quantity or price A_t with $A^t = (A_t, A_{t+1}, \dots)$. For a collection of plans to form an equilibrium, they must satisfy the following five conditions.

- 1) Given $\hat{K}_t, \hat{S}_t, \hat{B}_t, \hat{V}_t, R^t, H^t$, and W^t , the borrower's plans for $\hat{C}^t, \hat{S}^t, \hat{X}^t, \hat{K}^t, \hat{I}^t, \hat{B}^t$, and \hat{V}^t
 - are consistent with the initial given values of $\hat{K}_t, \hat{S}_t, \hat{B}_t$, and \hat{V}_t ;
 - obey the rules for accumulating market capital, home capital, and collateral value;
 - satisfy the borrower's borrowing and budget constraints in every quarter; and
 - yield a higher utility value for the borrower than any other plans that satisfy this condition's other requirements.
- 2) Given $\tilde{K}_t, \tilde{S}_t, \tilde{B}_t, \tilde{V}_t, R^t, H^t$, and W^t , the saver's plans for $\tilde{C}^t, \tilde{S}^t, \tilde{X}^t, \tilde{K}^t, \tilde{I}^t, \tilde{B}^t$, and \tilde{V}^t
 - are consistent with the initial given values of $\tilde{K}_t, \tilde{S}_t, \tilde{B}_t$, and \tilde{V}_t ;
 - obey the rules for accumulating market capital, home capital, and collateral value;

- satisfy the saver's borrowing and budget constraints in every quarter; and
 - yield a higher utility value for the saver than any other plans that satisfy this condition's other requirements.
- 3) For all $\tau > 0$, $\hat{B}_{t+\tau} + \tilde{B}_{t+\tau} = 0$.
 - 4) For all $\tau > 0$, $K_{t+\tau} = \hat{K}_{t+\tau} + \tilde{K}_{t+\tau}$.
 - 5) For all $\tau > 0$, $K_{t+\tau}$ and N are the capital and labor choices that maximize the representative firm's profits given H_t and W_t .

The first two conditions just require each of the households to do the best they can (measured with their utility values) with what they have got. The third condition states that the net supply of risk-free bonds in the economy equals zero. Thus, if one household wishes to borrow, the other must lend. The fourth condition says that the economy's stock of market capital must equal the sum of the two households' market capital stocks. And the final condition requires the rental rate of capital and the wage rate to induce the profit-maximizing representative firm to rent the entire available capital stock and employ all of the available hours of work.

investing the proceeds in bonds, and consuming the principal and interest in year $t + 1$. By its construction, this experiment leaves consumption in all years after $t + 1$ unchanged. If Δ is small, then the utility loss in year t is Δ / \tilde{C} and the discounted utility gain in year $t + 1$ equals $\tilde{\beta}R\Delta / \tilde{C}$. Here, R is the steady-state interest rate. Since the original consumption plan maximized utility, this slight change cannot increase utility. The change also cannot lower utility because if it did, then the analogous experiment that increases consumption in year t by borrowing Δ and repaying it in year $t + 1$ would increase utility. Therefore, we have that:

$$\frac{\Delta}{\tilde{C}} = \tilde{\beta}R \frac{\Delta}{\tilde{C}}.$$

Eliminating common terms from both sides yields our first important result, $R = 1/\tilde{\beta}$. That is, the saver's discount rate determines the steady-state rate of interest alone. Credit market regulation that changes either π and ϕ has no long-run effect on the interest rate.

Since the borrower is less patient than the saver, the experiment of borrowing Δ in year t and paying it

back at the interest rate $1/\tilde{\beta}$ in year $t + 1$ would increase her utility. However, the collateral requirements prevent her from doing this. Since the borrower exhausts her borrowing opportunities in the steady state, we can calculate her leverage ratio as:

$$-\frac{\hat{B}}{\hat{S}} = \frac{(1-\pi)\delta}{\phi}.$$

Thus, increasing either π or ϕ directly reduces the borrower's leverage ratio in the long run. Since the saver *purchases* bonds, we set his leverage ratio to zero.

Quantitative analysis of increasing equity requirements

Although the steady-state analysis reveals that equity requirements have no long-run effect on interest rates, it does not rule out substantial short-run effects in the wake of a reform. Investigating this possibility requires a quantitative analysis of the model's equilibrium, which we provide here. For this, we assign values to the model's parameters that reflect the equity

TABLE 1								
Calibrated parameter values								
Equity requirement	π	ϕ	α	λ	δ	$\tilde{\beta}$	$\hat{\beta}$	θ
High	0.16	0.0315						
			0.3	0.025	0.01	$\frac{1}{1.01}$	$\frac{1}{1.015}$	0.37
Low	0.11	0.0186						

Note: See the text for further details.
Source: Campbell and Hercowitz (2009).

requirements of household debt typical of the late 1990s and early 2000s. After calculating the model's steady state with these values, we raise the equity requirement parameters to values more typical of the period before the financial deregulation of the early 1980s. We then calculate the model's equilibrium paths for all quantities and prices when households start with the capital and debt stocks from the initial steady state (associated with low equity requirements) but face the new higher equity requirement parameters. In the long run, the economy's interest rate, its capital stocks, and the debt owed by the borrower to the saver converge to their values in the steady state calculated with the new parameters. We focus on the model economy's transition from the initial steady state to the other steady state following the parameter change.

Table 1 lists the parameter values we use for this experiment. All of them are taken from our earlier analysis of credit market deregulation in Campbell and Hercowitz (2009). We consider two configurations for the equity requirement parameters: high and low. In both cases, π equals the average of typical down payments on homes and automobiles weighted by their shares of durable purchases, and ϕ equals the average repayment rates of home mortgages and automobile loans weighted by their shares of household debt. The parameters for the high regime were chosen using observations of household debt and loan terms from before the financial liberalization of 1983, while the choice of the low regime's parameters used similar observations from 1995 through 2001. The required down payment for a home capital good equals 16 percent of its value in the high regime and 11 percent in the low regime. The model's remaining parameters are held constant across the two regimes. Campbell and Hercowitz (2009) provide justification for the specific values chosen. We note here only that the choice of $\tilde{\beta}$ produces an annual steady-state interest rate of 4.02 percent.

For our experiment, we start the economy at the model's steady state calculated with the parameters

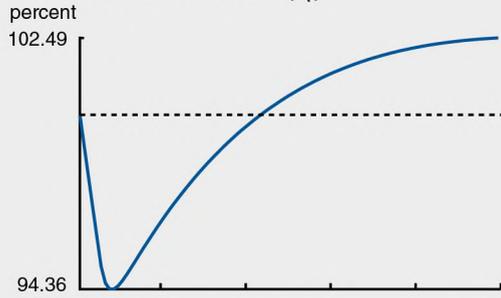
from the low regime. We suppose that, without warning, the parameters switch to those of the high regime. Both of the model's households expect the change to be permanent. Given the initial values of $\hat{S}_t, \hat{V}_t, \hat{B}_t, \hat{B}_t^i, \hat{S}_t^i,$ and \hat{K}_t^i , from the steady state associated with low equity requirements, we calculate the model's equilibrium. Figure 3 contains plots of the resulting equilibrium paths for the model's key variables. Panels A, B, C, and D plot the values of both households' consumption choices, and panels E and F display the evolution of the productive capital stock and the wage. All of these have been scaled so that their values in the original steady state equal 100 percent. Panel G shows the interest rate in annual percentage points, and panel H shows the household leverage ratio in percentage points.

In the model, there are two reasons for the borrower to purchase durable goods: They create a flow of valuable services, and they enable the expansion of debt. The re-regulation of household debt markets reduces the size of this second incentive, and so the reform initially makes the borrower wish to reduce her stock of durable goods. Indeed, the borrower purchases *no* durable goods for six quarters following the re-regulation (figure 3, panel A). This decline in durable purchases together with the acceleration of principal repayment required by the higher value of ϕ reduces loan demand, so both the interest rate and the household leverage ratio fall as expected. The leverage ratio starts at 38.17 percent, falls rapidly while the borrower purchases no durable goods, and then declines more gradually toward its new long-run level of 23.37 percent (figure 3, panel H). The interest rate falls rapidly from its initial value of 4.02 percent to its trough three years after re-regulation, 3.24 percent—a decrease of 78 basis points (figure 3, panel G). Thereafter, the interest rate rises very slowly towards its steady-state value. Even 25 years after re-regulation, the interest rate is 36 basis points below its original value. Apparently, it takes a long time indeed to reach the long run.

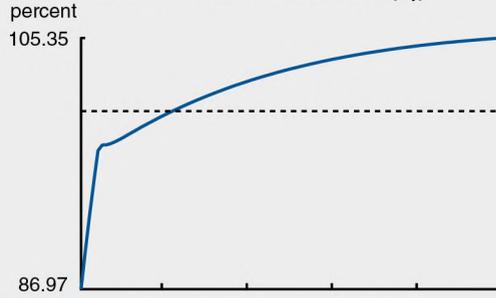
FIGURE 3

The model's equilibrium following financial re-regulation

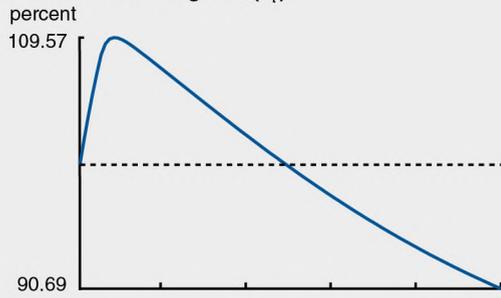
A. Borrower's durable goods (\hat{S}_t)



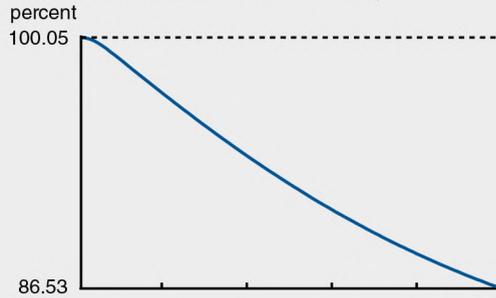
B. Borrower's nondurable consumption (\hat{C}_t)



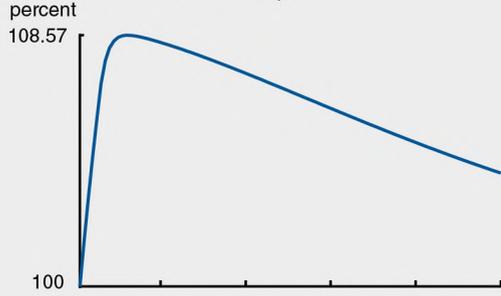
C. Saver's durable goods (\tilde{S}_t)



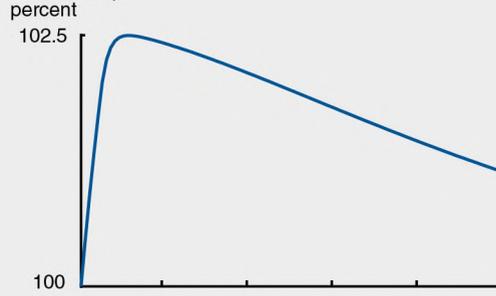
D. Saver's nondurable consumption (\tilde{C}_t)



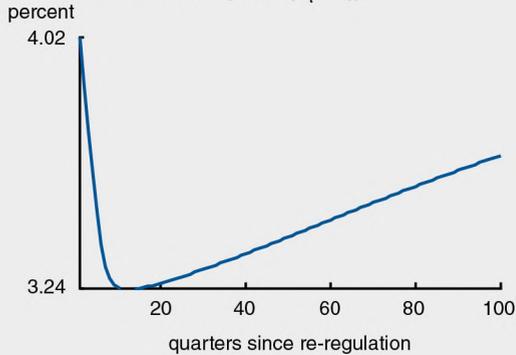
E. Stock of market capital (K_t)



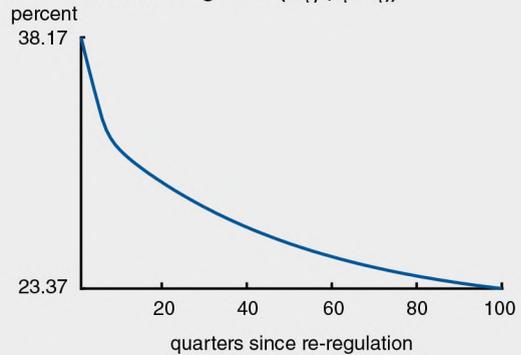
F. Wage (W_t)



G. Annual interest rate ($400 \times (R_t - 1)$)



H. Household leverage ratio ($-\hat{B}_t / (\tilde{S}_t + \hat{S}_t)$)



Notes: Panels A through F indicate the variable relative to its value in the initial steady state, which has been set to equal 100 percent (the dashed horizontal line). The values on the vertical axis in each panel are the variable's minimum and maximum values attained in the 100 quarters following re-regulation.

A note on welfare

In this article, we have examined interest rates in the wake of the deregulation and re-regulation of financial markets. Appropriate monetary policy requires understanding and forecasting persistent interest rate changes, so our results can contribute to that discussion. However, for those who set financial market policy, the interest rate serves only as a means to an end. Policymakers instead concern themselves with how adopting a given policy changes the welfare of borrowers and savers. In the model, we can measure welfare with the two households' utility values after the policy change. Comparing these with the analogous utility values from the pre-reform steady state provides the desired welfare assessment.

Before reporting on the actual welfare changes, it is worth returning to figure 3. It shows that after 25 years, the saver consumes much less of both goods than he did before the reform (panels C and D). At the same time, the borrower consumes more of both goods (panels A and B). Although the economy has not yet reached its new steady state in that time, these changes also characterize the long run. Therefore, the reform unambiguously increases the borrower's welfare while decreasing the saver's.

The long-run welfare changes are only tangentially interesting for policymakers; they care about the total welfare change that accounts for the short-run transition from one steady state to another. In the short run, the borrower's consumption of both goods falls (figure 3, panels A and B). The saver's nondurable consumption slowly trends down (panel D). The saver's durable purchases rise to peak at about 10 percent above their pre-reform level, and then fall to their new steady state value (panel C).

In principle, the borrower's short-run utility loss could dominate her welfare calculation. This would be intuitive because re-regulation imposes a constraint on her decisions. The actual utility changes reported in table 2 show that this is not the case. The utility values themselves have no meaningful units, so all of the table's entries give the *permanent* percentage change in the consumption of both goods (starting from the original steady state) required to make the household's utility equal to its post-reform value.

In table 2, the first row reports the results for the experiment plotted in figure 3. The borrower's welfare change equals that from permanently and instantly raising her consumption of both durable and nondurable goods by 0.9 percent. The borrower is better off, even though she faces tighter constraints on her borrowing.

TABLE 2		
Consumption-equivalent welfare changes		
	Borrower	Saver
Baseline experiment (<i>percent</i>)	0.9	-8.4
Fixed <i>K</i> (<i>percent</i>)	-0.1	-4.8
Note: See the text for further details.		

This would be impossible if the interest rate she pays on her debts and the wage she receives for her labor were held constant. Of course, both of these variables also change in the short run, and the changes are favorable to the borrower: The interest rate falls, and the wage rises. These two are actually tightly connected. The interest rate decline increases the capital employed by the representative firm, which in turn raises wages. Put differently, the re-regulation induces the saver to invest more in productive capital and thereby benefit the borrower indirectly with higher wages.⁷

To determine whether the "direct" effect of lower interest rates or the "indirect" effect of higher wages contributes more to the borrower's welfare gain, we have run an experiment with the model in which we hold the stock of market capital fixed at its original steady-state level. Put differently, we force the saver to replace depreciated market capital and do not allow any further investment. In this experiment, the interest rate falls 129 basis points (to 2.73 percent) over three years; the wage remains constant by construction. In table 2, the row for fixed *K* reports the consumption equivalent welfare changes analogous to those from the previous experiment. Even though the fall in interest rates is much larger than before, the borrower's welfare gain becomes a loss. The change also cuts the saver's welfare loss substantially. Apparently, the indirect effects of financial re-regulation on consumer welfare can easily dominate its more easily envisioned direct effects.⁸

Since tightening consumer lending standards helps one household at the expense of the other, it is impossible to unambiguously state that such a policy change helps or hurts "society as a whole." A policymaker who cares only for the borrower would prefer tighter lending standards, while one who represents the saver's interests would be against them. A policymaker who wishes to keep both households' considerations in mind can come to either conclusion depending on the weights she assigns to the two households' preferences. We have been silent regarding how many "real" households the borrower and saver each represent because that detail is actually irrelevant for the model's equilibrium. As long as no single household thinks that it can

influence the wage or interest rate, nothing changes if we divide either household into 10, 100, or 1,000 smaller (but identical) households.

Nevertheless, the number of “actual” borrowers and savers clearly matters for a policymaker’s welfare calculations. In our favored interpretation of the model, the saver represents the 5 percent or 10 percent of households with the highest wealth, and the borrower represents the remainder. If 5 percent of households are savers, then tightening lending standards increases the average utility value of all households. However, the same tightening decreases average utility if 10 percent of households are savers. Therefore, we have little concrete advice to give a policymaker who wishes to base her judgment on changes in average utility. That is, we can identify winners and losers from tightening lending standards, but assessing whether or not this improves society lies well beyond our capabilities.

Conclusion

Empirically, times of expanding home leverage have had higher-than-average interest rates. Interest rates in the United States during the post-Korean War surge in household leverage were about 60 basis

points higher than their average in the period immediately after the leverage ratio had peaked. Similarly, interest rates fell about 200 basis points when the second sustained increase in household leverage ratios ended in 1995 (recall our discussion of figures 1 and 2). Of course, macroeconomic events other than changes in credit market regulation substantially influence interest rates. Nevertheless, these results give a range within which reasonable model predictions for the interest rate effects of financial re-regulation should fall. In the baseline version of our model in which the saver accumulates market capital, the interest rate falls 78 basis points over three years after financial re-regulation. Thereafter, the interest rate rises very slowly back to its original level. If we instead suppose that the stock of market capital is fixed and cannot be augmented, the analogous decline is about 130 basis points. These two specifications embody two extreme assumptions on the costs of adjusting market capital: none and infinite. Accordingly, we argue that any persistent decline in interest rates between 78 basis points and 130 basis points is a reasonable forecast in the wake of financial re-regulation.

NOTES

¹See Semer et al. (1986) and Olney (1991) for more information about household credit markets before the Great Depression.

²Green and Wachter (2005) provide a history of the spread of amortized mortgages in the United States.

³See Friedman (1992) for a discussion of government deficits and interest rates in the 1980s. Campbell and Hercowitz (2009) argue that rising demand for credit must have contributed to that decade’s high interest rates because otherwise household indebtedness would have declined as government deficits increased interest rates.

⁴Becker (1980) describes this long-run behavior of household indebtedness in detail.

⁵For an alternative view, see Kiyotaki (1998). He discusses one environment of *limited commitment* in which the creditors require down payments because collateral loses value upon repossession.

⁶To calculate $\hat{U}(\hat{C}^t, \hat{S}^t)$, choose a large integer τ and artificially set $\hat{U}(\hat{C}^{t+\tau}, \hat{S}^{t+\tau})$ to zero. Next, use the equation to calculate $\hat{U}(\hat{C}^{t+\tau-1}, \hat{S}^{t+\tau-1}), \hat{U}(\hat{C}^{t+\tau-2}, \hat{S}^{t+\tau-2}), \dots, \hat{U}(\hat{C}^t, \hat{S}^t)$. This calculation is obviously incorrect because the assumption upon which it is predicated is false. However, the error will generally be proportional to $\hat{\beta}^\tau$, which gets very small as τ becomes larger.

⁷It is important to note here that the borrower’s welfare increase does *not* reflect a paternalistic assumption built into the model that regulators can make better financial decisions than individual borrowers. Instead, it reflects the benefits accruing to all borrowers from them simultaneously reducing their loan demand. In this sense, financial re-regulation has the same effects as would the formation of a borrowers’ cartel to limit the demand for loans. All of the borrowers are made better off if they stick to the cartel agreement, but each one of them would like to deviate and expand her indebtedness so long as the others conform.

⁸In this experiment, the saver’s welfare improves when his choices over market capital are restricted. Just as before with the borrower’s welfare following financial re-regulation, this welfare improvement can be interpreted as a cartelization of savers. If all savers commit to not increasing market capital, they can all avoid paying higher wages on the transition path. This increases their welfare, even though it further reduces the interest rate. Of course, each individual saver would like to expand his purchases of market capital if all other savers stick to the cartel agreement.

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Measuring the equilibrium real interest rate

Alejandro Justiniano and Giorgio E. Primiceri

Introduction and summary

In conducting monetary policy, policymakers find it useful to monitor the performance of the economy relative to some benchmark. For instance, the policy decision whether to raise or lower the short-term nominal interest rate might be affected by the deviations of current inflation from policymakers' comfort zone, of output from potential output, and of the real interest rate (current nominal rate minus expected future inflation) from its equilibrium value (the rate that would be consistent with output at its potential level). Unfortunately, these benchmark concepts are not directly observed in the data, but can only be defined in the context of a specific theoretical framework.

Over the past decade, the *new Keynesian model* has become the workhorse for the analysis of monetary policy. This model departs from the neoclassical framework of the 1980s by assuming imperfect competition in goods and labor markets and “sticky” (meaning rigid or inflexible) prices and wages—neoclassical models assume prices and wages are flexible and adjust quickly. These ingredients in the new Keynesian model alter the transmission of fundamental shocks perturbing the economy and allow monetary policy to have temporary real effects.

The *equilibrium real interest rate* is a crucial concept in the new Keynesian class of models. This rate represents the real rate of return required to keep the economy's output equal to *potential output*, which, in turn, is the level of output consistent with flexible prices and wages and constant markups in goods and labor markets (Woodford, 2003; and Galí, 2008).¹ Meanwhile, the difference between the *ex ante real interest rate*—the nominal interest rate minus expected inflation—and the equilibrium real interest rate is defined as the *real interest rate gap*.

In the new Keynesian model, the real interest rate (RIR hereafter) gap is central to the determination of

output and inflation. Loosely speaking, if this RIR gap is positive, output will decline relative to potential. This is because people will be inclined to postpone spending decisions today to take advantage of higher returns to savings. All else being equal, a negative output gap will then put downward pressures on prices and wages because of weaker aggregate demand. Conversely, a negative RIR gap will typically be associated with a positive output gap, setting in motion inflationary forces—higher demand leads to higher prices.

The main policy implication of this observation is that policymakers concerned with maintaining output close to its potential level should set short-term nominal interest rates—the policy instrument of most central banks—in order to minimize the RIR gap. In the absence of a trade-off between stabilizing inflation and output, this simple policy prescription would also completely stabilize inflation. In practice, however, there may well be a trade-off between the two objectives of output and inflation stabilization.² Nonetheless, the equilibrium RIR constitutes a natural benchmark for the conduct of monetary policy, and the RIR gap can be viewed as providing some indication of the stance of monetary policy (Neiss and Nelson, 2003).

While the equilibrium RIR is theoretically appealing, its use in guiding monetary policy decisions faces at least two major hurdles. First and foremost, the equilibrium RIR is not directly observable in the data, limiting its usefulness as a target for monetary policy in practice.³ Moreover, rather than being constant, the

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equilibrium RIR fluctuates over time in response to a variety of shocks to preferences and technology that perturb the economy.

Second, setting nominal interest rates to track the equilibrium RIR may not be feasible at times because of the existence of the *zero bound*; that is, nominal interest rates cannot be set lower than zero. Indeed, the equilibrium RIR may fall enough to induce a positive RIR gap, even with the nominal interest rate at zero. Output would then decline below potential, engendering deflation. In this way, the gap helps us to gauge the constraint imposed by the zero bound on monetary policy. With short-term nominal interest rates now at historically low levels in the United States and a number of other industrialized countries, this scenario is receiving a lot of attention from both the academic community and policymakers.

Given the importance that the equilibrium RIR plays for the design of monetary policy in modern macroeconomic models, our purpose in this article is to provide an estimate of this unobservable variable. We do so by inferring it from an empirical new Keynesian model fitted to U.S. quarterly data on a few key macroeconomic variables from 1962:Q1 through 2008:Q4.⁴

Specifically, our analysis accomplishes three objectives. First, we describe the historical evolution of the equilibrium RIR. We find that this rate has been negative at times, particularly in the late 1970s and, most interestingly, during the latest recession.

Second, we estimate the short-term RIR gap as the difference between the current (as opposed to future) *ex ante* RIR and the equilibrium RIR. This provides some indication of the stance of monetary policy. Consistent with the anecdotal view, the estimated short-term RIR gap suggests that policy was loose during most of the 1970s. In contrast, policy would seem to have been tight at the end of our sample. However, this mostly reflects the zero bound problem—policymakers' inability to lower short-term nominal interest rates below zero—and provides a rationale for the nonconventional policy measures undertaken by the Federal Reserve during the most recent recession, such as direct purchases of longer-term securities and the creation of special facilities and programs (for example, the Term Asset-Backed Securities Loan Facility, or TALF) intended to increase access to credit.

Finally, we compare the evolution of the short-term and long-term RIR gaps, where the latter is defined as the sum of the current and expected future short-term RIR gaps or, alternatively, the difference between the *ex ante* long-term RIR and the equilibrium long-term RIR. Long-term rates reflect the path of

current and expected future short-term rates. Therefore, long-term gaps summarize private expectations about future macroeconomic outcomes and monetary policy, providing a more forward-looking measure of the policy stance. For instance, according to this measure, policy was not loose in the 2002–06 period, which preceded the recent economic downturn. This characterization of the policy stance contrasts with what is suggested by the short-term RIR gap and, in particular, with the view of several commentators (see, for instance, Taylor, 2007).

Several papers have tackled the estimation of the equilibrium RIR before, most notably Laubach and Williams (2003) and Kozicki and Clark (2005). In contrast to these earlier studies, our estimate of the equilibrium RIR is based on a *micro-founded* model, which builds on the optimizing behavior of households and firms seeking to maximize their utility and profits. In this respect, this article is related to the approach of Neiss and Nelson (2003), Amisano and Tristani (2008), and, in particular, Edge, Kiley, and Laforge (2008). However, in contrast to these earlier studies, we stress the importance of both current and expected future RIR gaps for the determination of macroeconomic outcomes.

As with all empirical work based on structural models, our results may be sensitive to some aspects of the model specification. To illustrate this point, we compare our results across two models that differ in scale, shocks, and transmission mechanisms of these disturbances.

The article is organized as follows. First, we provide a brief description of our baseline model economy. Then, we describe the data and the estimation approach. Next, we present the main results—that is, we present our estimates of the equilibrium RIR and RIR gaps. We also discuss the robustness of these estimates when inferred from a larger-scale model. We conclude with a few comments and caveats to our analysis, particularly with regard to the current economic situation. More specifically, we note how the larger-scale model also suggests the presence of positive short-term and long-term RIR gaps for the fourth quarter of 2008. This provides a further rationale for the Federal Reserve's response to the current crisis with nonconventional measures to ease monetary policy. We do, however, emphasize the need to enhance these models' ability to capture the interplay between the financial sector and the real economy, particularly in light of recent events.

The model

In this section, we sketch our baseline new Keynesian model and analyze two of its key equilibrium relations—the aggregate demand and supply

equations. The presentation is mostly narrative, with most of the technical details relegated to the appendix. Interested readers can refer to Justiniano and Primiceri (2008) for greater details on the model, or they can see the comprehensive treatment of new Keynesian models in Woodford (2003) and Galí (2008), as well as the excellent primer by Galí and Gertler (2007). For simplicity, relative to Justiniano and Primiceri (2008), the model here abstracts from the roles of habit formation, indexation, and endogenous capital accumulation. We present the results based on a larger-scale model with these additional features as a robustness check in a later section.

There are five types of agents in our model economy: 1) households, 2) employment agencies, 3) firms producing intermediate goods, 4) firms producing final goods, and 5) the monetary authority. We now briefly describe the behavior of each of them.

Households

We assume that we have a large number of households seeking to maximize their stream of current and expected future utility, which depends positively on their consumption of a single final good and negatively on the number of hours they work for the production of intermediate goods. Each household is the sole supplier of a specialized type of labor that it sells to the employment agencies in exchange for wages. Rather than taking wages as given—as under the neoclassical assumption of perfect competition—each household has some market power and can post its wage. This, in turn, determines the amount of their specialized labor demanded by the employment agencies.

We introduce sticky wages in the labor market by assuming that at each point in time only a random fraction of households can change their posted wage. Hence, when setting its wage, each household takes into consideration not only current but also future labor demand and costs of working. For example, if future labor demand is expected to rise, households will preemptively post higher wages, since they might not be able to do so in the near future.

Finally, all households have access to savings through two types of assets: one-period government bonds and state-contingent securities, which pay only if a certain future state is realized. The former are used to smooth consumption over time. State-contingent securities serve instead to insure against the idiosyncratic risk arising from the uncertainty about the length of time before households will be able to reset their wages.

Employment agencies

Employment agencies mediate the demand and supply of labor between households and firms

producing intermediate goods. Their role is to purchase all types of specialized labor supplied by households and bundle them into a single homogenous labor input sold to intermediate goods firms. Employment agencies operate in a perfectly competitive market, taking the wage received for the labor bundle as given and making zero profits.

Intermediate goods producers

A large number of intermediate goods producers combine technology with labor inputs purchased from employment agencies to produce differentiated intermediate goods, which are then sold to final goods producers. Each of the intermediate goods producers has some market power and can therefore post the price of its good. This, in turn, determines the amount of its output demanded by the final goods producers.

We introduce sticky prices in the goods market by assuming that at each point in time only a random fraction of firms can change their posted price. Hence, when setting its price, each firm takes into consideration not only current but also future demand and marginal costs, where the latter depend on wages. For example, if future demand is expected to rise, producers will preemptively increase prices, since they might not be able to adjust them in the near future.

Final goods producers

Final goods producers mediate between intermediate goods producers and households. They produce the final good by bundling all intermediate goods into a single final homogenous commodity purchased by households. Final goods firms maximize profits as well, but in contrast to the intermediate goods producers, they operate under perfect competition, taking the price for the final good as given and making zero profits.

Monetary authority

The central bank determines monetary policy by setting the short-term nominal interest rate in response to price inflation and output growth. This interest rate rule is a variant of the instrument rule proposed by Taylor (1993), the Taylor rule, which approximates the historical behavior of the U.S. federal funds rate. According to this rule, nominal interest rates rise more than one-to-one with inflation and fall in response to output contractions.

Demand, supply, and the equilibrium RIR

Before presenting our estimation results, we highlight the main insights of the two crucial equilibrium relations in the model. This helps explain the roles of the equilibrium RIR and RIR gaps in the determination of output and inflation.

Aggregate demand

In the model, aggregate spending is determined by the behavior of the representative household, which seeks to smooth consumption over time by investing its savings in one-period government bonds. This optimizing behavior results in the following (log-linearized) *aggregate demand* equation, which is also known as the *IS* equation:

$$1) \quad \hat{y}_t = E_t \hat{y}_{t+1} - \hat{r}_t,$$

where y_t and r_t are output and the RIR, respectively, and the hat symbol ($\hat{\cdot}$) denotes deviations from the equilibrium level. Hence, \hat{y}_t denotes the output gap, and \hat{r}_t stands for the short-term RIR gap. Intuitively, according to the aggregate demand equation, fluctuations in the short-term RIR gap induce deviations of the output gap from its expected future value, $E_t \hat{y}_{t+1}$, where the operator E_t denotes households' expectation of future values conditional on the information available today.

Equation 1 can be iterated forward to express the output gap today only as a function of the current and expected future short-term RIR gaps. This procedure yields the expression

$$2) \quad \hat{y}_t = - \sum_{j=0}^{+\infty} E_t \hat{r}_{t+j},$$

by which the output gap is negatively associated with the long-term RIR gap. The latter corresponds to the sum of current and expected future short-term RIR gaps.⁵ Notice, therefore, that if the long-run RIR gap is negative, the output gap will be positive, and vice versa.

Aggregate supply

In terms of the supply side, intermediate goods firms set prices according to the current and expected future evolution of marginal costs and demand conditions. Profit-maximizing behavior results in the following (log-linearized) *aggregate supply* or *Phillips curve* equation:

$$3) \quad \pi_t = \beta E_t \pi_{t+1} + \kappa s_t + \lambda_{\pi,t},$$

where π_t and s_t stand for price inflation and real marginal costs, respectively, and $\lambda_{\pi,t}$ is a markup shock that represents exogenous variation to the level of markup desired by intermediate goods producers. Finally, β is a constant very close to one that represents the temporal discount factor, and κ is a positive constant

that is inversely related to the degree of price stickiness. Intuitively, inflation exceeds its expected future level either if real marginal costs increase or if intermediate goods firms change their desired markup of prices over marginal costs for other reasons exogenous to the model.

To highlight the importance of the RIR gap for inflation determination, we briefly analyze a special case of our model obtained by assuming perfectly flexible wages. Under this assumption, real marginal costs are proportional to the output gap. Hence, all else being equal, a positive output gap will cause inflation to rise relative to its expected future level. Moreover, if the output gap is projected to remain positive in the future, expected future inflation will also increase, further fueling the rise in current inflation. That is, current and expected future RIR gaps engender pressures on prices through their effects on aggregate demand. This crucial insight also holds in our general model with wage rigidities, although with sticky wages the link between the output gap and real marginal costs is more complex.

RIR gaps and monetary policy

Equations 1 and 3 highlight the importance of RIR gaps for output and inflation determination. Current and future expected deviations of ex ante RIRs from their corresponding equilibrium values affect the output gap, which, in turn, influences the inflation rate. Since the ex ante RIRs depend on the nominal interest rates set by the monetary authority, the conduct of monetary policy is central to the behavior of the RIR gaps and, hence, output and inflation.

Consider, for instance, a central bank that seeks to stabilize price inflation and the output gap. Absent any markup shocks ($\lambda_{\pi,t}$), the central bank can achieve full stabilization of both output and inflation by committing to set nominal interest rates according to an appropriate instrument rule that delivers a zero RIR gap at every point in time.

However, as we mentioned in the introduction, tracking the equilibrium RIR may not be feasible when the zero bound on nominal interest rates becomes binding. Put another way, sometimes the equilibrium RIR may fall enough that, even with the short-term nominal interest rate at zero, positive RIR gaps would emerge. In this case, according to the model, output would decline relative to potential and inflation would fall.

Even abstracting from the zero bound, in practice optimal monetary policy is more involved than the simple prescription of tracking the equilibrium RIR. This is due to the fact that markup shocks bring about a trade-off between stabilizing the output gap and inflation.⁶ Nonetheless, despite these considerations,

the equilibrium RIR remains an important reference point for the conduct of monetary policy, assuming that it can be accurately estimated and forecasted. This is the task we undertake next.

Model solution and estimation

In this section, we provide a brief overview of the approach that we adopt to estimate the model's parameters and to infer the evolution of the latent (unobservable) variables. The discussion is somewhat technical, although we do not aim to provide a comprehensive overview of the techniques we used. For more details on these techniques, interested readers should refer to An and Schorfheide (2007).

Model solution and state-space representation

The model we described in the preceding section has a solution of the form

$$4) \quad \tilde{\xi}_t = G(\theta)\tilde{\xi}_{t-1} + M(\theta)\tilde{\varepsilon}_t,$$

where the state vector $\tilde{\xi}_t$ collects all variables except for the shocks. The elements of $\tilde{\xi}_t$ are expressed in (log) deviations from the model's nonstochastic steady state, which corresponds to the constant values of all variables that the economy would converge to in the absence of shocks. The shocks inducing temporary deviations from the steady state are stacked in the vector $\tilde{\varepsilon}_t$. Meanwhile, $G(\theta)$ and $M(\theta)$ are matrices whose elements are functions of the vector of model structural parameters, denoted by θ . Our goal is to estimate these parameters and to uncover the historical behavior of the unobserved variables in the state vector.

In fact, while some elements of the state vector are directly observed in the data (for instance, inflation and output), others are not (such as the equilibrium RIR and expected inflation). Therefore, in order to estimate the model, equation 4 must be combined with an additional set of equations specifying which elements of the state vector are observed in the data.

The general form of this additional set of equations is

$$5) \quad x_t = Z(\tilde{\xi}_t + C(\theta)),$$

where Z is a matrix mapping the elements of $\tilde{\xi}_t$ into x_t (the vector of observable data) and where C is a vector of constant terms (which may depend on θ) representing the steady state of the observable elements of $\tilde{\xi}_t$. Equations 4 and 5 constitute the transition and measurement equations of a linear *state-space* model.

Data

We estimate the model, using five series of U.S. quarterly data: 1) real per capita gross domestic product (GDP), 2) per capita hours worked, 3) real per capita wages, 4) quarterly inflation, and 5) the short-term nominal interest rate. We construct real GDP by dividing nominal GDP by the population aged 22–65 and the GDP deflator.⁷ For hours, we use a measure of hours in all sectors of the economy following Francis and Ramey (2008). This is also our source for the population series. Real wages correspond to nominal compensation of employees from the U.S. Bureau of Economic Analysis's national income and product accounts (NIPAs), divided by hours and the GDP deflator; for the nominal interest rates, we use the effective federal funds rate. The sample period spans 1962:Q1 through 2008:Q4.⁸ We do not de-mean or de-trend any series.

Bayesian inference

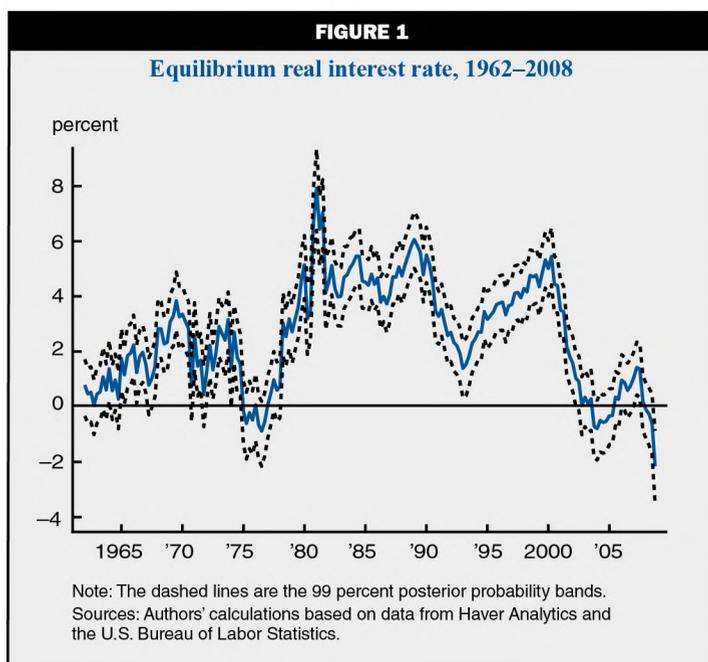
The state-space representation of the model allows us to use a very powerful algorithm known as the Kalman filter to estimate the parameters θ and retrieve the most likely path of the unobservable elements of $\tilde{\xi}_t$. We discuss each in turn.

A natural way to estimate the model is to find the value of the parameters θ that maximizes the *likelihood* function. The likelihood function summarizes all information about θ contained in a sample of data and plays a pivotal role in econometrics and statistics. The likelihood function of our state-space model can be evaluated using the Kalman filter.

In practice, however, the likelihood function associated with most modern macroeconomic models is typically a complicated nonlinear function of the model parameters. This makes finding a unique value that maximizes the likelihood a rather arduous task. For this reason, most of the recent literature estimating macro models has turned to Bayesian methods, which discipline the set of plausible values for θ through the use of prior information.

Bayesian inference then seeks to characterize the distribution of θ that results from combining the likelihood function with the prior information. This is known as the *posterior* distribution, from which we can compute the location of a parameter (mean or median) and a measure of uncertainty. For instance, the uncertainty surrounding θ can be conveyed by reporting posterior probability bands that contain the range of values that parameters are likely to take with, say, 99 percent probability.

Prior beliefs about the elements of θ may be informed by theory or simply reflect and summarize



information not contained in the estimation sample. In practice, this prior information is formulated by specifying a certain distribution for each element of the parameter vector, centered at a particular value (mean) and with an associated measure of uncertainty (standard deviation).

Once we have estimated the model's parameters, we can employ the Kalman filter to sequentially and systematically update our guess for the unobserved elements of the state vector. More precisely, at each point in time, our guess for $\tilde{\xi}_t$, based on data available in the previous quarter, is updated after we observe the data for the current period. This *filtered* (or one-sided) estimate for the state vector forms the basis for our guess on the value of the state vector next period, which we also update once we have data for the next quarter, and so on.

Having followed this procedure for all periods, we can go back and revise the filtered estimate of $\tilde{\xi}_t$, conditional not only on information up to time t but also on the whole sample of data. We call the state vector emerging from this procedure the *smoothed* (or two-sided) estimate. We analyze these estimates in the next section.

Equilibrium RIR and RIR gaps in the estimated model

We do not report the estimated parameters in this article. They are similar to those of Justiniano and Primiceri (2008), who use a longer sample. Here, we

focus on our estimates of the equilibrium RIR and the RIR gaps.

The equilibrium RIR

Figure 1 plots the smoothed estimate of the equilibrium RIR (solid blue line). It is also important to characterize the uncertainty surrounding the estimated equilibrium RIR, particularly since this is cited as a possible concern regarding its usefulness for monetary policy analysis. Therefore, we also report uncertainty bands (dashed black lines), which represent the values this variable is likely to have taken with 99 percent probability. We first highlight a few properties of the smoothed estimate and later discuss these probability bands.

The first thing to notice is that the inferred equilibrium RIR has fluctuated substantially over our sample, with a standard deviation of 1.94 percent around a mean of 2.6 percent (annualized).⁹

A second interesting feature of figure 1 is that the equilibrium RIR has turned negative in a few instances. This occurred around 1975 and the end of 2008—two recession dates, as determined by the National Bureau of Economic Research—and during the 2003–04 period. These episodes were characterized by a substantial decline in the federal funds rate in response to weak economic conditions. However, the 2008 episode is the only one in our sample for which the uncertainty bands are completely below zero.

Indeed, the third interesting observation is that the equilibrium RIR has plummeted in the latest part of the sample. In particular, during the latest recession, the equilibrium RIR seems to have recorded by far its largest decline, with an estimate for 2008:Q4 of roughly -2.15 percent.

The tightness of the posterior probability bands deserves some comment. In particular, the precision with which the equilibrium RIR is estimated perhaps seems implausible, especially considering that these bands account for the uncertainty surrounding both the unobserved states and the model parameters. It is important to keep in mind, however, that these probability bands abstract from model uncertainty. That is, alternative specifications of the model (for example, a different historical characterization of U.S. monetary policy or a model with additional propagation mechanisms and/or shocks) might deliver different estimates of the equilibrium RIR. We return to this issue in the section explaining the larger-scale model.

This being said, the cross-sectional dispersion at different points in time is larger than perhaps suggested visually by figure 1. For example, figure 2 plots the posterior distribution of the equilibrium RIR for the last point in the sample, 2008:Q4. Values of the equilibrium RIR are on the horizontal axis, with the vertical line drawn at the median of -2.15 percent, which coincides with the estimate reported in the previous figure. Notice that this distribution has a range from roughly -4 percent to -0.5 percent, with hardly any weight assigned to values close to zero. Therefore, our model-based estimates suggest that it is quite likely that the equilibrium RIR became negative in 2008. To what extent did this induce positive RIR gaps? We address this key issue next.

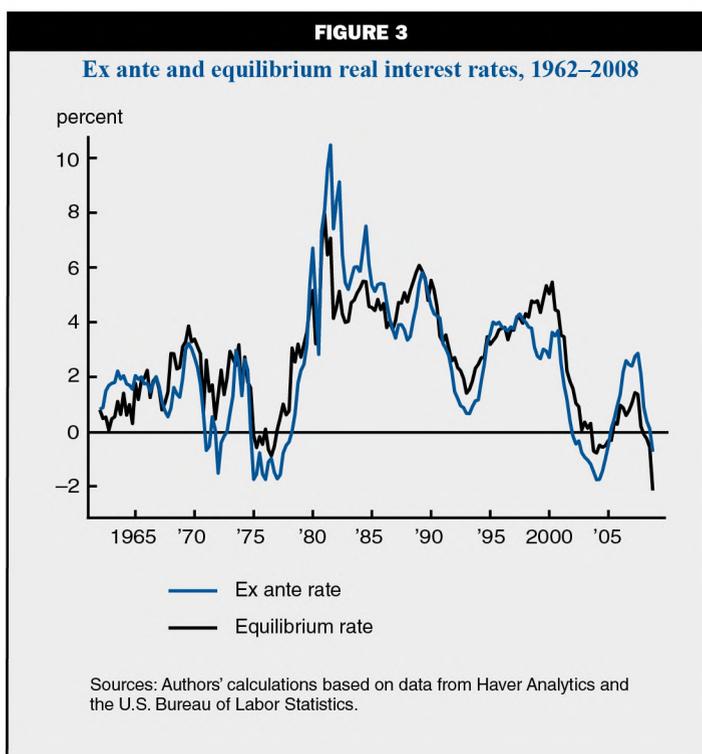
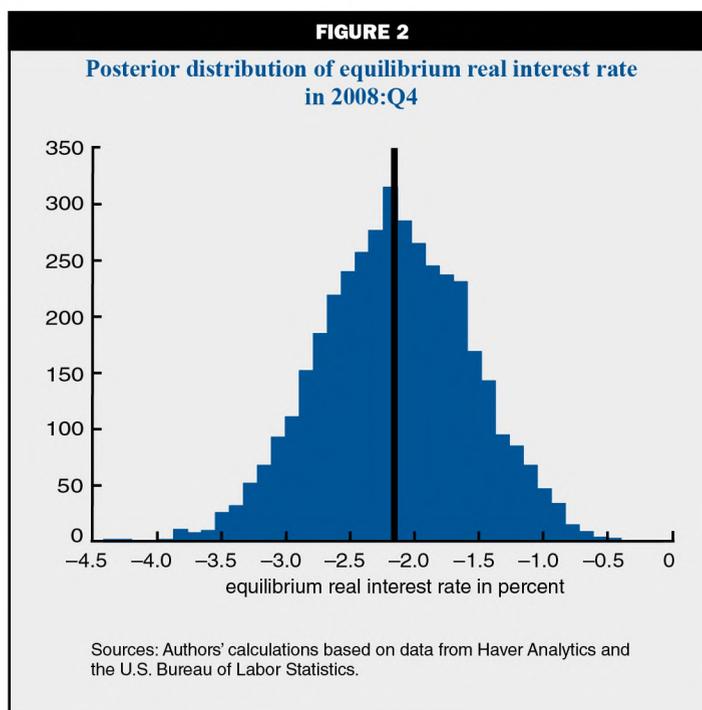
The short-term RIR gap

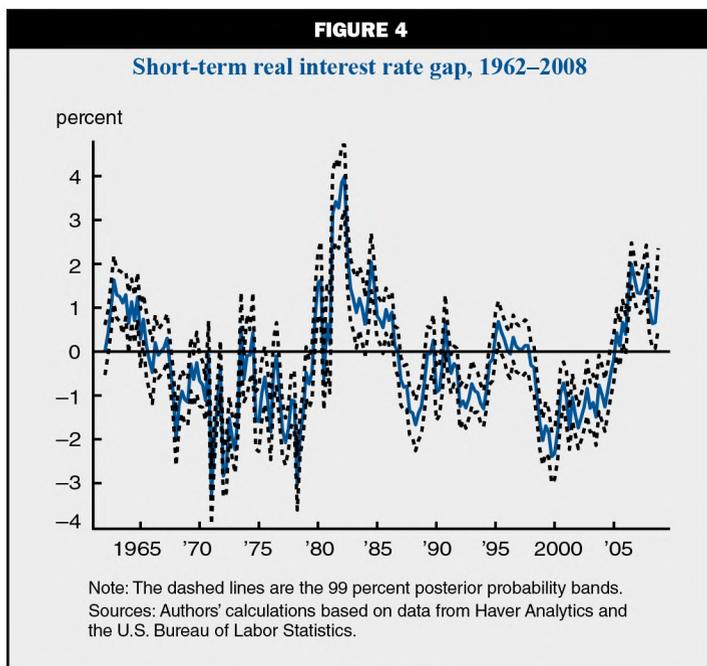
The ex ante RIR is given by the difference between the nominal interest rate and the inflation rate expected for next quarter. While the former is directly observable in our data, the latter is part of the unobservable state vector and must be backed out using the Kalman filter.

Figure 3 shows the smoothed estimate of the ex ante RIR (blue line) together with the equilibrium RIR (black line). The mean of the ex ante RIR is 2.37 percent (annualized) with a standard deviation of 2.45 percent. These statistics are similar to those corresponding to the equilibrium RIR. The overall contours of these two series coincide, although they have differed at times.

In order to highlight the discrepancies between the ex ante RIR and the equilibrium RIR, figure 4 plots their difference together with its 99 percent probability bands. We refer to this difference as the short-term RIR gap, in order to distinguish it from the long-term gap that we analyze next. Note that the short-term gap has also fluctuated considerably over time, with an average of -0.33 percent and a standard deviation of 1.28 percent.

As we noted earlier, the short-term RIR gap is commonly taken as a measure of the monetary policy stance. And indeed,





at least for some episodes, the evolution of the RIR gap aligns well with the anecdotal characterization of monetary policy that we see in the literature. For instance, according to our estimates, the equilibrium RIR exceeded the ex ante real interest rate during most of the 1970s, exactly when U.S. inflation was at historically high levels. This is consistent with the view that monetary policy during this period was characterized by an insufficient response to the rise in inflation (Clarida, Galí, and Gertler, 2000). Similarly, the significant increase in the short-term RIR gap in the early 1980s accords well with the conventional view that the disinflation in the U.S. economy was engineered by a substantial policy tightening under then-Federal Reserve Chairman Paul Volcker.

The long-term RIR gap

While the behavior of the short-term RIR gap presented in figure 4 squares quite well with the conventional view, there are a few caveats that call for caution in interpreting this gap as a good proxy for the stance of monetary policy. In particular, as we explained earlier, it is important to recognize that the whole path of expected future short-term RIR gaps—rather than just its contemporaneous value—matters for the determination of output and inflation in the new Keynesian model (see equation 2, p. 17). From this perspective, we might judge the monetary policy stance better by looking at the long-term RIR gap, which summarizes the information contained in the current and expected future values of the federal funds rate, inflation, and the equilibrium RIR.

To this end, figure 5 compares the short-term RIR gap (blue line) with the evolution of the long-term one (black line). Although the two series often move together—the correlation coefficient is equal to 0.56—the message about the stance of monetary policy implied by the two lines differs during some historical episodes.

The 2002–06 period provides an interesting example. In 2002:Q3 the federal funds rate stood at 1.75 percent, but it had declined to 1 percent by 2003:Q3, and remained there for the next three quarters. The federal funds rate then rose gradually, reaching 5.25 percent in 2006:Q3. Some have argued that monetary policy was too accommodative during this period (for example, Taylor, 2007). Although the negative value of the short-term RIR gap seems to accord with this claim (blue line), the positive value of the long-term RIR gap (black line) does not support the

view that policy was too expansionary. In particular, it suggests that the private sector expected a decline of the equilibrium RIR or a monetary tightening.

The difference between short-term and long-term gaps toward the end of the sample is also informative. For instance, our estimate of the short-term RIR gap in 2008:Q4 is roughly 1.5 percent. This suggests that, according to the model, the federal funds rate of 0.5 percent was probably above the equilibrium RIR. Furthermore, it suggests that the zero bound on nominal interest rates would have been binding before additional interest rate cuts could have closed the short-term RIR gap. In addition, the estimated long-term RIR gap exceeds 3 percent. Taken at face value, this would suggest that at the end of 2008, positive short-term gaps were expected to persist and the zero bound was expected to bind beyond a single quarter.

Before we interpret this result as indicative of contractionary monetary policy, we must acknowledge that these gaps can only reflect the stance of *conventional* monetary policy. By this we mean the Federal Reserve's management of the short-term nominal interest rate. During the current economic crisis, the Federal Reserve has also employed a variety of nonconventional policy measures; and these measures have been reflected in the changing size and composition of the Federal Reserve's balance sheet. Our simple analysis suggests that these measures have been appropriate, insofar as both the short-term RIR and long-term RIR exceeded the equilibrium

RIR. However, these extraordinary measures are not reflected in our analysis of the short-term and long-term RIR gaps.

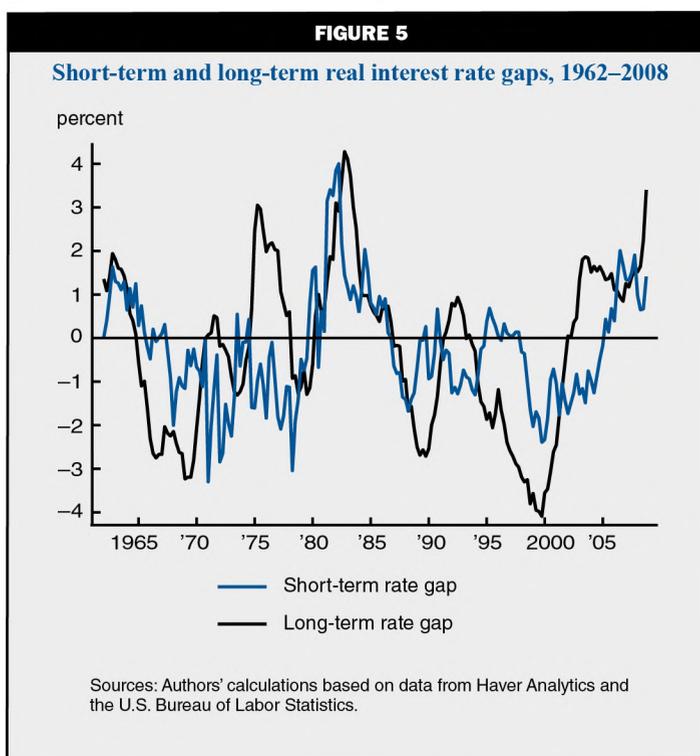
A larger-scale model

The baseline model can be summarized in a few simple equations that, as discussed, clearly highlight the role of the equilibrium RIR for the dynamics of output and inflation. This simplicity, however, comes at the expense of abstracting from other features that impart more realism to the model. In particular, additional shocks can be included and other mechanisms added (such as endogenous capital accumulation) through which disturbances influence the evolution of the economy. For this reason, we test the robustness of our main conclusions by using a larger-scale model estimated on a richer data set. This extended model is discussed in Justiniano and Primiceri (2008) and is based on the well-known studies of Christiano, Eichenbaum, and Evans (2005) and Smets and Wouters (2007).

Relative to our baseline model, the larger-scale model includes the additional propagation mechanisms provided by endogenous capital accumulation, investment adjustment costs, a choice of capital utilization, habit formation in consumption, and indexation in both prices and wages. These features are essentially meant to increase the length of time for which a given shock will affect the evolution of the economy. There are three additional disturbances perturbing the model economy, specifically, shocks to the marginal efficiency of investment, to the disutility of labor, and to government spending. Finally, we estimate the model over the same sample, 1962:Q1 through 2008:Q4, but we incorporate additional data on consumption and investment.

Figure 6 reports the smoothed estimates of the equilibrium RIR and the ex ante RIR, as well as the short-term and long-term RIR gaps. In each panel, the black line reproduces the estimates from the baseline model and the blue line corresponds to estimates from the extended model.

Panel A highlights the fact that the cyclical pattern of the equilibrium RIR is very similar across models, although the equilibrium RIR is substantially more volatile in the larger-scale model.¹⁰ One implication of this finding is that, according to the extended model, the equilibrium RIR has declined below zero more



frequently than what is predicted by our baseline framework. Furthermore, the decline in the current downturn, while substantial, is not as dramatic by historical standards as suggested by the baseline model.

Since the inferred ex ante RIR (panel B) is almost identical across models, it is not surprising that the short-term RIR gap (panel C) and long-term RIR gap (panel D) are more volatile in the larger-scale model as well. Notice also that the estimates from our baseline model and larger-scale model co-move more closely in the case of the long-term gap, for which the two lines essentially overlap during the latest part of the sample.

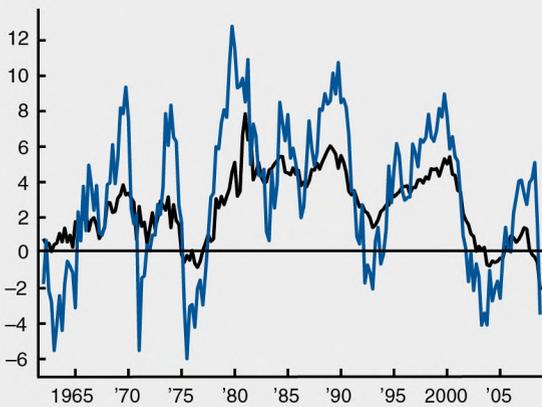
Regarding the 2002–06 period, the discrepancy between the short-term and long-term RIR gaps is far less evident in the larger-scale model than in our baseline model. However, both measures in the larger-scale model remain positive or very close to zero. This confirms our earlier observation that policy may not have been as accommodative during this period as has been suggested (for example, Taylor, 2007).

Consistent with the baseline model, the larger-scale framework also predicts large positive short-term and long-term RIR gaps for the fourth quarter of 2008. However, the same caveats we raised earlier about interpreting these endpoint estimates as reflecting the policy stance apply to the larger-scale model as well.

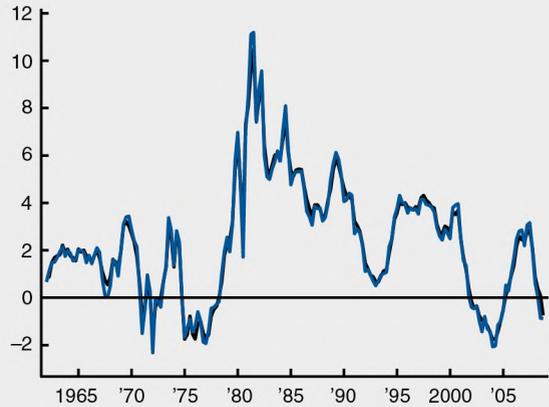
FIGURE 6

Real interest rate levels and gaps in the baseline model and larger-scale model, 1962–2008

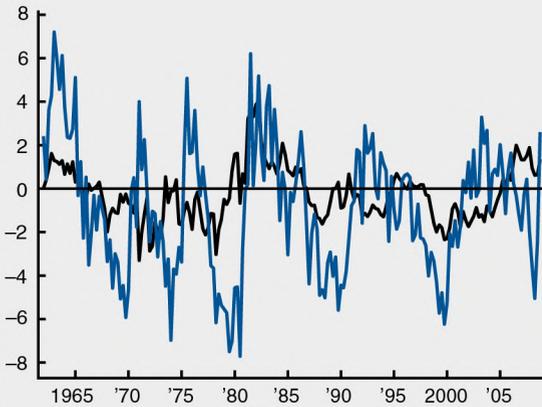
A. Equilibrium real interest rate
percent



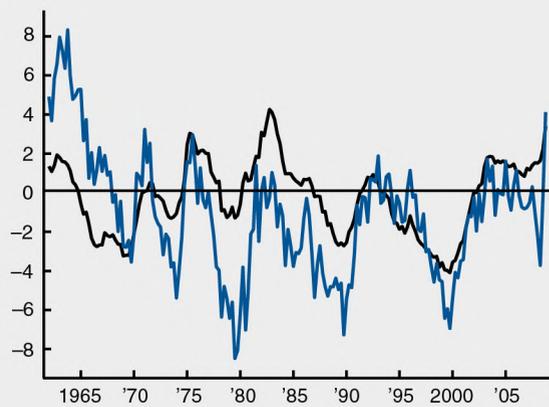
B. Ex ante real interest rate
percent



C. Short-term real interest rate gap
percent



D. Long-term real interest rate gap
percent



— Baseline model
— Larger-scale model

Sources: Authors' calculations based on data from Haver Analytics and the U.S. Bureau of Labor Statistics.

Overall, despite some obvious discrepancies, we view these results as an important assessment of robustness of our main findings. Furthermore, they suggest—in line with our earlier hypothesis—that model uncertainty is likely to be a crucial factor surrounding the measurement of the unobservable equilibrium RIR and related components. This source of uncertainty is sometimes ignored in studies presenting model-based estimates of the RIR, although our findings suggest that this should be a major issue for further empirical work in this area.

Conclusion

In this article, we study the evolution of the equilibrium RIR and RIR gaps, using both a prototypical new Keynesian model and a larger-scale model similar to those in Christiano, Eichenbaum, and Evans (2005) and Smets and Wouters (2007). Our estimates point to a substantial degree of time variation in the equilibrium RIR. Moreover, we find that this rate has sometimes become negative in the post-war period. In particular, our analysis suggests that the equilibrium RIR fell sharply below zero toward the end of 2008 (although

the magnitude of this decline relative to historical standards is model dependent), resulting in positive short-term and long-term expected RIR gaps. This provides some support for the Federal Reserve's response to the current crisis with nonconventional measures to ease monetary policy.

We conclude by noting that the models we use here, even the larger-scale one, are to some extent

very stylized and have some shortcomings. One of these shortcomings is the absence of an explicit theoretical framework of the financial sector and financial frictions. It would be useful to analyze how the introduction of these additional features would affect our results (as, for instance, in Christiano, Motto, and Rostagno, 2007). These features seem particularly relevant for the analysis of current economic events.

NOTES

¹Hence, we could alternatively refer to the equilibrium real interest rate as the real interest rate at potential. We prefer the former terminology because it is more popular in the literature and policy discussions, as exemplified by the discussion in Ferguson (2004). Meanwhile, potential output is proportional, but lower than the efficient level of output. The efficient level of output is the level of output under perfect competition and, therefore, with zero markups. In the goods market, the markup is defined as the amount by which prices exceed the marginal cost of production. In the labor market, the markup is defined as the excess of wages over the marginal cost of supplying labor.

²Exogenous variations in desired markups, usually referred to as markup shocks, introduce such a trade-off (see, for example, Clarida, Galí, and Gertler, 1999).

³Potential output is not directly observable either, and the policy implications of its measurement have received substantial attention following the work of Orphanides (2001). See also Justiniano and Primiceri (2008).

⁴We also estimate the model's unknown parameters and subsequently extract all unobserved model-based variables, such as expected inflation next period.

⁵While seemingly daunting to compute, the long-run rates can be backed out from the Lagrange multiplier of the household's budget constraint.

⁶If wages are rigid, optimal monetary policy must attribute some weight to wage inflation stabilization as well.

⁷All data except for hours are from Haver Analytics. We are very grateful to Shawn Sprague, of the U.S. Bureau of Labor Statistics, for providing us the series of hours in all sectors of the economy.

⁸We use the eight years prior to the sample period to initialize the Kalman filter.

⁹This result is consistent with the large degree of time variation reported by Laubach and Williams (2003) and Edge, Kiley, and Laforge (2008), but stands in contrast to the analysis of Neiss and Nelson (2003), who argue that the equilibrium real interest rate exhibits very little volatility.

¹⁰The main reason the equilibrium RIR in the larger-scale model is more volatile is that this model includes habit formation.

APPENDIX: MODEL EQUATIONS

We present the main equations of the model for each of the five classes of agents described in the main text.

Households

The expected discounted stream of utility that each household j maximizes is given by

$$A1) \quad E_t \sum_{s=0}^{\infty} \beta^s b_{t+s} \left[\log C_{t+s} - \varphi \frac{L_{t+s}(j)^{1+\nu}}{1+\nu} \right],$$

where C_t denotes consumption, and the second argument of the utility function represents the marginal disutility of each household's specific labor, $L(j)$, that depends on the parameter ν , known as the inverse Frisch elasticity of labor supply. Future utility is discounted at the rate β , and b_t is a "discount factor" shock affecting both the marginal utility of consumption and the marginal disutility of labor. The logarithm of b_t is modeled as a Gaussian autoregressive process of order 1, denoted as AR(1) for short.

At every point in time t , each household's sources and uses of income must be equal, as summarized by the budget constraint

$$P_t C_t + T_t + B_t \leq R_{t-1} B_{t-1} + Q_{t-1}(j) + \Pi_t + W_t(j) L_t(j),$$

where T_t is lump-sum taxes and transfers, B_t denotes holdings of government bonds, R_t is the gross nominal interest rate, $Q_t(j)$ is the net cash flow from participating in state-contingent securities that insure against idiosyncratic risk, and Π_t is the per capita profit that households get from owning the intermediate goods firms.

Following Erceg, Henderson, and Levin (2000), we permit in every period only a fraction $1 - \xi_w$ of households to reset their wages to minimize the expected discounted stream of labor disutility for the periods in which the posted wage is anticipated to remain in place,

$$E_t \sum_{s=0}^{\infty} \xi_w^s \beta^s b_{t+s} \left\{ -\varphi \frac{L_{t+s}(j)^{1+\nu}}{1+\nu} \right\}.$$

This is subject to the labor demand function of employment agencies specified next. Wages for the remaining ξ_w fraction of households are indexed to steady-state inflation and productivity.

Employment agencies

Competitive employment agencies operate in competitive markets and bundle each household's specialized labor $L_t(j)$ into a homogenous labor input according to

$$L_t = \left[\int_0^1 L_t(j)^{\frac{1}{1+\Lambda_{w,t}}} dj \right]^{1+\Lambda_{w,t}}.$$

Homogeneous labor is sold to intermediate goods firms. Profit maximization and the zero profit condition imply a specialized labor demand function,

$$L_t(j) = \left(\frac{W_t(j)}{W_t} \right)^{-\frac{1+\Lambda_{w,t}}{\Lambda_{w,t}}} L_t,$$

where $W_t(j)$ is the wage paid by the employment agencies to the household supplying labor of type j , and W_t is the hourly wage paid by intermediate goods firms for their homogenous labor input. The demand schedule for labor j is decreasing in the relative wage and depends on the elasticity of substitution among varieties of labor given by $\Lambda_{w,t}$. Notice that this elasticity is time varying, and we assume that $\log(1 + \Lambda_{w,t})$ is a Gaussian independent and identically distributed (i.i.d.) process. In the literature this is referred to as the wage markup shock, and it is analyzed in detail in Justiniano and Primiceri (2008).

Intermediate goods producers

A monopolistically competitive firm produces the intermediate good $Y_t(i)$ with the production function

$$Y_t(i) = A_t L_t(i)^\alpha,$$

where $L_t(i)$ denotes the bundled labor input purchased from employment agencies for the production of good i , and A_t represents a productivity shock. We model A_t as nonstationary, with its growth rate following a Gaussian AR(1) process.

As in Calvo (1983), at each point in time a fraction ξ_p of firms cannot reoptimize their prices and index them to steady-state inflation. The remaining fraction $1 - \xi_p$ of firms post a new price $\tilde{P}_t(i)$ to maximize the expected discounted stream of profits for the periods in which the new price is anticipated to remain in place,

$$E_t \sum_{s=0}^{\infty} \xi_p^s \beta^s \Lambda_{t+s} \left\{ \left[\tilde{P}_t(i) \pi^s \right] Y_{t+s}(i) - W_{t+s} L_{t+s}(i) \right\},$$

where Λ_{t+s} is the marginal utility of consumption used to value future income, subject to the goods demand function specified in the next section.

Final goods producers

Perfectly competitive firms produce the final good Y_t by bundling all intermediate goods according to

$$Y_t = \left[\int_0^1 Y_t(i)^{\frac{1}{1+\Lambda_{p,t}}} di \right]^{1+\Lambda_{p,t}}.$$

Profit maximization and zero profit condition for the final goods producers imply the following demand function for the intermediate good i :

$$Y_t(i) = \left(\frac{P_t(i)}{P_t} \right)^{-\frac{1+\Lambda_{p,t}}{\Lambda_{p,t}}} Y_t,$$

where P_t corresponds to the aggregate price level. The demand schedule for intermediate good i is decreasing in its relative price, and depends on the elasticity of

substitution $\Lambda_{p,t}$ among varieties of intermediate goods. This elasticity is time varying, and we assume that $\log(1 + \Lambda_{p,t})$ is a Gaussian i.i.d. process. This disturbance is known as the price markup shock.

Monetary authority

The Taylor type rule for the short-term nominal interest rate, R_t , is given by

$$\frac{R_t}{R} = \left(\frac{R_{t-1}}{R} \right)^{\rho_R} \left[\left(\frac{\prod_{s=0}^3 \pi_{t-s}}{\pi_t^*} \right)^{1/4} \right]^{\phi_\pi} \left[\left(\frac{Y_t / Y_{t-4}}{e^\gamma} \right)^{1/4} \right]^{\phi_Y} e^{\varepsilon_{R,t}},$$

with R being the steady state for the gross nominal interest rate and $\varepsilon_{R,t}$ being a Gaussian i.i.d. monetary policy shock. The parameters ϕ_π and ϕ_Y capture how aggressively the monetary authority responds to variations in inflation and output growth over the current and previous three quarters. There is a time-varying inflation target π_t^* , which evolves exogenously according to a Gaussian AR(1) process. Finally, notice that short-term nominal interest rates are adjusted gradually, as given by ρ_R , referred to as the smoothing coefficient.

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What is behind the rise in long-term unemployment?

Daniel Aaronson, Bhashkar Mazumder, and Shani Schechter

Introduction and summary

As we entered 2010, the average length of an ongoing spell of unemployment in the United States was more than 30 weeks—the longest recorded in the post-World War II era. Remarkably, more than 4 percent of the labor force (that is, over 40 percent of those unemployed) were out of work for more than 26 weeks—we consider these workers to be long-term unemployed. In contrast, the last time unemployment reached 10 percent in the United States, in the early 1980s, the share of the labor force that was long-term unemployed peaked at 2.6 percent. Although there has been a secular rise in long-term unemployment over the last few decades, the sharp increases that occurred during 2009 appear to be outside of historical norms. Further, this trend may present important implications for the aggregate economy and for macroeconomic policy going forward.

The private cost of losing a job can be sizable. In the short run, lost income is only partly offset by unemployment insurance (UI), making it difficult for some households to manage their financial obligations during spells of unemployment (Gruber, 1997; and Chetty, 2008). In the long run, permanent earnings losses can be large, particularly for those workers who have invested time and resources in acquiring knowledge and skills that are specific to their old job or industry (Jacobson, LaLonde, and Sullivan, 1993; Neal, 1995; Fallick, 1996; and Couch and Placzek, 2010). Health consequences can be severe (Sullivan and von Wachter, 2009). Research even suggests that job loss can lead to negative outcomes among the children of the unemployed (Oreopoulos, Page, and Stevens, 2008) and to an increase in crime (Fougère, Kramarz, and Pouget, 2009).

All of these costs are likely exacerbated as unemployment spells lengthen. The probability of finding a job declines as the length of unemployment increases. Although there is some debate as to exactly

what this association reflects, it is certainly plausible that when individuals are out of work longer, their labor market prospects are diminished through lost job skills, depleted job networks, or stigma associated with a long spell of unemployment (Blanchard and Diamond, 1994).¹ For risk-averse households that cannot insure completely against a fall in consumption as they deplete their precautionary savings, the welfare consequences of job loss rise as unemployment duration increases. Welfare implications are particularly severe during periods of high unemployment for individuals with little wealth (Krusell et al., 2008).

In this article, we analyze the factors behind the recent unprecedented rise in long-term unemployment and explain what this rise might imply for the economy going forward. Using individual-level data from the U.S. Bureau of Labor Statistics' *Current Population Survey* (CPS), we show that all of the substantial rise in the average duration of unemployment between the mid-1980s and mid-2000s can be explained by demographic changes in the labor force, namely, the aging of the population and the increased labor force attachment of women (Abraham and Shimer, 2002). But only one-half of the increase in average duration of unemployment at the end of 2009 relative to that of the early 1980s may be due to demographic factors. This suggests that other factors have come into play more recently.

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In particular, we attribute the sharp increase in unemployment duration in 2009 to especially weak labor demand, as reflected in a low rate of transition out of unemployment into employment, and a smaller portion of this increase (perhaps 10 percent to 25 percent) to extensions in unemployment insurance benefits.²

We show that, in any given month, individuals with longer unemployment spells are less likely to be employed the following month. This suggests that the average ongoing spell of unemployment is likely to remain longer than usual well into the economic recovery and expansion, plausibly keeping the unemployment rate above levels observed in past recoveries. For example, we find that if the current distribution of unemployment duration resembled historical distributions, the unemployment rate would be roughly 0.4 percentage points lower than it is today. Nevertheless, we find no evidence that high levels of long-term unemployment will have a sizable impact on compensation growth going forward.

We begin by presenting some descriptive facts about trends and business cycle movements of unemployment duration. We then analyze how much of the increase in the recent average duration of unemployment compared with that of the previous severe recession and its aftermath (in 1982–83) can be explained by changes in the demographic, industrial, and occupational composition of the labor force versus changes in the average duration of unemployment within the various groups. We next consider how much of the remaining increase can be attributed to weak labor demand and extensions of unemployment benefits. Finally, we examine how high levels of long-term unemployment may affect the unemployment rate and compensation growth going forward.

The rise of long-term unemployment

We begin by reviewing some facts about unemployment spell length. Long-run estimates of unemployment duration are available back to the late 1940s from the *Current Population Survey*, a monthly survey of 60,000 or more households. Respondents are 16 years and older and are asked to classify themselves as employed, unemployed, or out of the labor force. Those unemployed are further asked how long, in weeks, their unemployment has lasted. As a result, the CPS duration measures are based on ongoing spells of unemployment and are not measures of completed spell length.

Figure 1 plots the average (and median) duration of unemployment from 1948 (and 1967) through the end of 2009. Over the past half century, the average length of spells of unemployment have increased, from 11.3 weeks in the 1960s to 11.8 weeks in the 1970s,

11.9 weeks in the 1980s, 15.0 weeks in the 1990s, and 17.4 weeks in the 2000s.³ Figure 2 plots the share of the unemployed that are short-term (fewer than five weeks) versus long-term (more than 26 weeks). There has been a pronounced shift over time in the composition of the unemployed by duration, with a particularly sharp change in 2009. Long-term unemployment accounted for 10 percent of the unemployed in the 1950s and 1960s; it reached 26 percent in the early 1980s; and it averaged roughly 20 percent between 2002 and 2007, but reached 40 percent as of December 2009.⁴ By the end of last year, over 4 percent of the labor force was long-term unemployed.

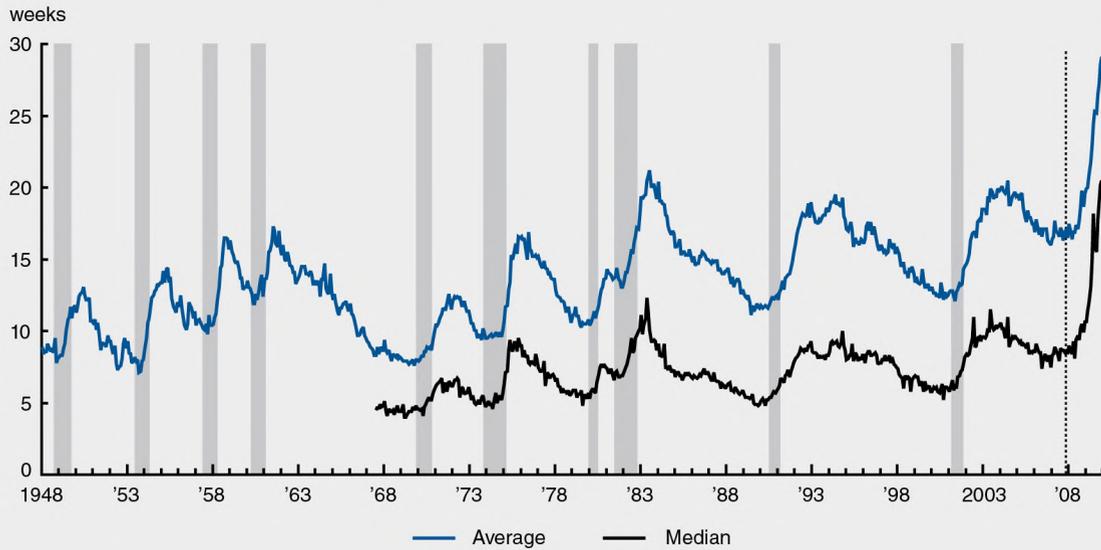
The average duration of unemployment is countercyclical—that is, it increases when the overall economy is shrinking, as figure 1 makes clear. Therefore, figure 3, panel A presents a scatter plot of average duration of unemployment against the unemployment rate to provide a simple way of comparing durations conditional on the unemployment rate. Each blue or black box represents a month. The black line represents the relationship between the unemployment rate and average duration of unemployment over the period 1948–2007. Because the line is upward sloping, it illustrates that worse labor market conditions (higher unemployment rates) are associated with longer unemployment spells. In particular, through 2007, an extra 1 percentage point on the unemployment rate was associated with spells that lasted 1.2 weeks longer on average.

For the most recent period, we use black boxes to represent months between December 2007 (the start of the most recent recession) and December 2009 in figure 3, panel A. Note that all the black boxes lie near the top of the cloud of blue boxes, highlighting that the average unemployment spell tends to be much longer now for any given unemployment rate. As the economy weakened and the unemployment rate rose, the length of unemployment spells increased—and at a pace that was fairly typical for a recession. This is represented by the black boxes that lie roughly parallel to the black line. But, starting in June 2009 (the half dozen or so black boxes on the right side of panel A), unemployment spells began to lengthen to unprecedented levels. Much of this spike in average duration of unemployment is driven by the unmistakable increase in the share of the unemployed out of work for more than 26 weeks, highlighted by the black boxes in figure 3, panel B. For instance, the average length of unemployment during the last six months of 2009 was over seven weeks longer than that of the first six months of 1983, when unemployment had peaked at 10.8 percent.

Looking forward, we should expect to see a historically long average duration of unemployment for

FIGURE 1

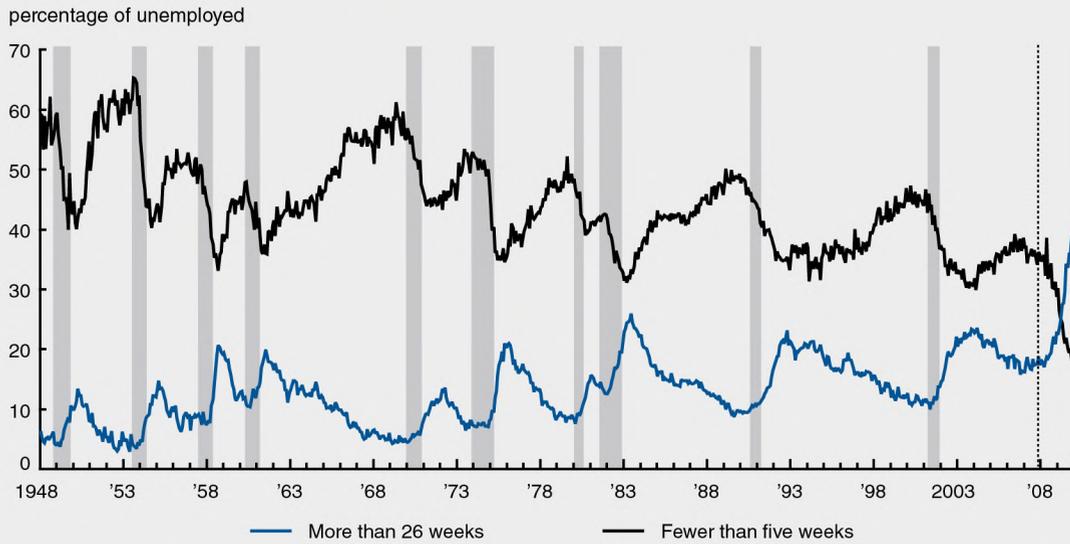
Average and median duration of unemployment, 1948–2009



Note: The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research; the dashed vertical line indicates the most recent business cycle peak.
Source: U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

FIGURE 2

Short-term and long-term share of unemployment, 1948–2009



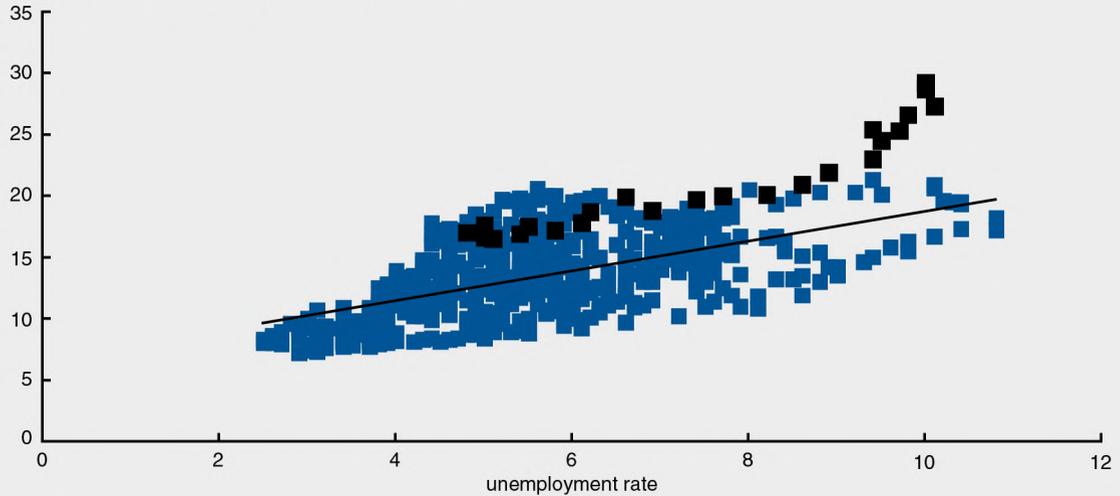
Note: The shaded areas indicate official periods of recession as identified by the National Bureau of Economic Research; the dashed vertical line indicates the most recent business cycle peak.
Source: U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

FIGURE 3

Unemployment rate, duration of unemployment, and long-term share of unemployment, 1948–2009

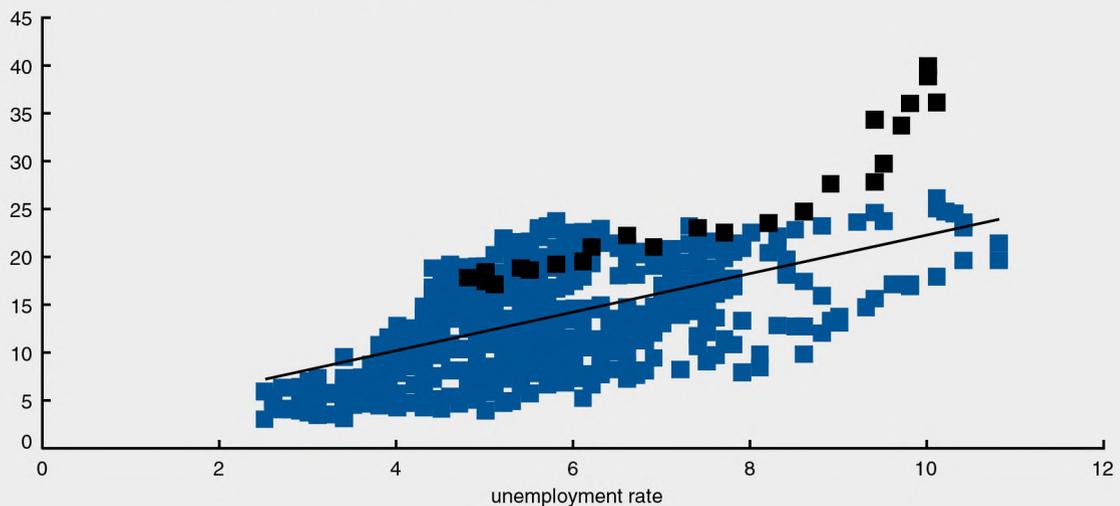
A. Average duration of unemployment versus unemployment rate

average duration of unemployment in weeks



B. Long-term share of unemployment versus unemployment rate

percentage of the unemployed who are long-term unemployed



■ January 1948–November 2007

■ December 2007–December 2009

Notes: In panel A, the black line represents the relationship between the unemployment rate and average duration of unemployment over the period January 1948–November 2007. In panel B, the black line represents the relationship between the unemployment rate and the share of the unemployed who are more than 26 weeks unemployed over the period January 1948–November 2007.

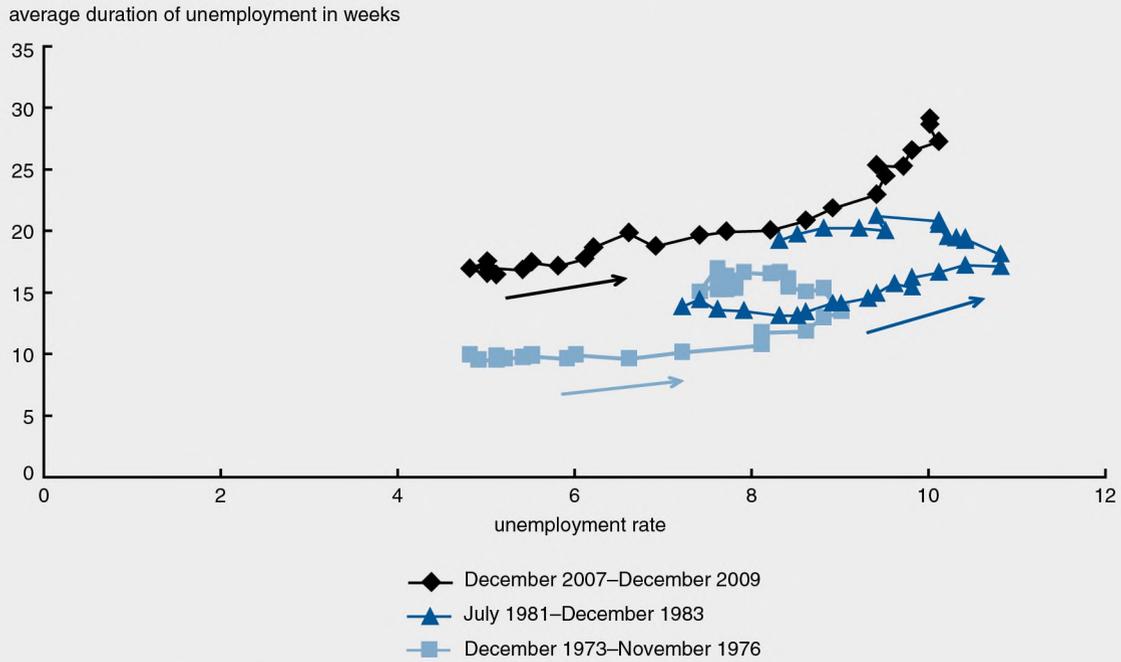
Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

some time, since it is typical for average spell length to rise well past the business cycle trough. This is apparent in figure 4, which plots the cyclical pattern in the average duration of unemployment versus the unemployment rate for several selected cycles. In both the mid-1970s and the early 1980s (blue lines), average

duration stayed persistently high, even as the unemployment rate began to decline.⁵ As labor demand picks up early in a recovery, employers might turn to unemployed workers with shorter spells first, leaving the unemployment pool increasingly composed of those with relatively longer spells. Sequential hiring

FIGURE 4

Unemployment rate versus average duration of unemployment for selected business cycles



Source: U.S. Bureau of Labor Statistics, *Current Population Survey*, from Haver Analytics.

patterns like this may be due in part to a selection effect: Those who are less employable are the ones who are likely to remain unemployed longer and are less likely to be rehired. However, the lower reemployment probability of the long-term unemployed may also be due to diminished job skills, weakened social networks, and the assumption by some employers of poor worker quality that accompany those with longer spells. Declines in job separations, which we discuss in more detail later, may also reduce the number of short spells of unemployment in the early stages of a recovery.

Unemployment duration versus other labor market measures

It is important to emphasize that the recent spike in the duration of unemployment not only is quite large by historical standards but also stands out relative to the recent deterioration in many other key labor market indicators, including three key measures used to gauge labor market slack: the unemployment rate, a broader unemployment rate (the U.S. Bureau of Labor Statistics' U-6 rate),⁶ and total payroll employment. That observation can best be seen from a very simple statistical model that uses gross domestic product (GDP) growth to generate out-of-sample forecasts of these labor market

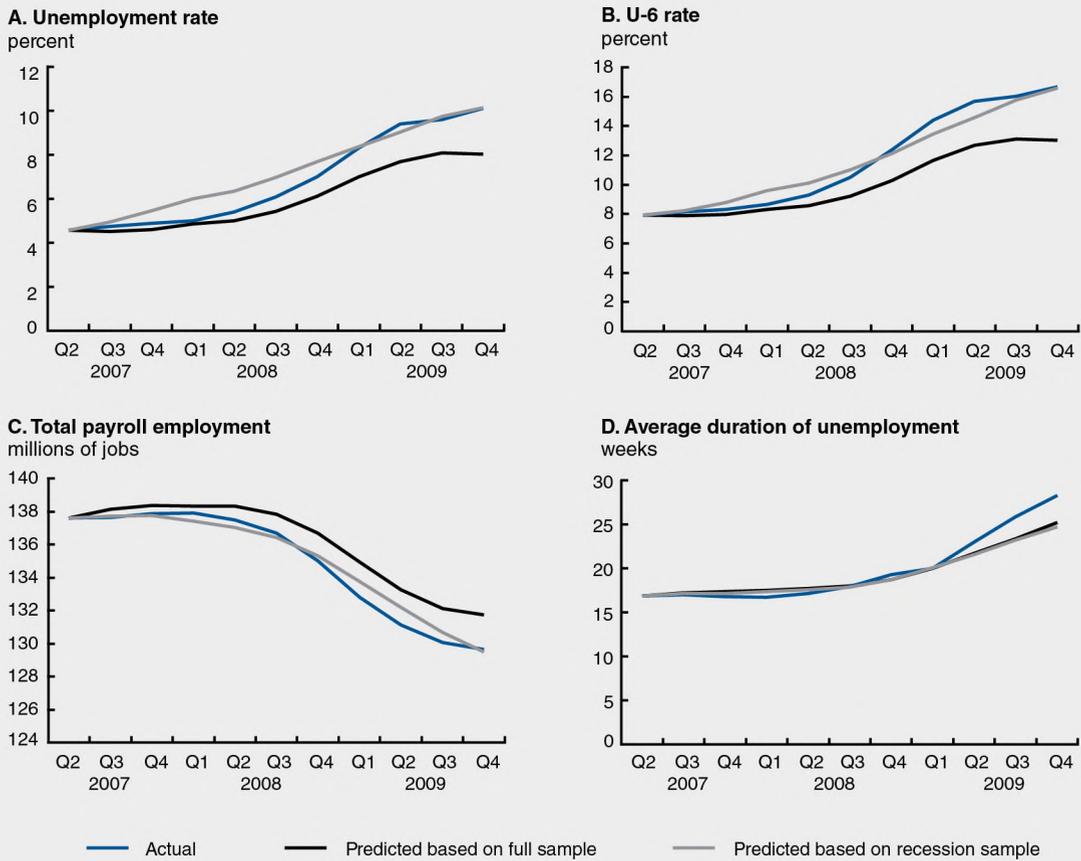
measures. This exercise when applied to the unemployment rate is the basis for what is often referred to as "Okun's law."⁷ We follow Aaronson, Brave and Schechter (2009) and use two samples to estimate these relationships: 1) all data from the first quarter of 1978 through the second quarter of 2007 and 2) data solely from the recessions during that period.

Figure 5 shows the results for four measures of the labor market—namely, the unemployment rate, the U-6 rate, total payroll employment, and the average duration of unemployment. Each panel of figure 5 contains three colored lines. The blue line represents the actual data, the black line is the forecast based on the data from our full sample, and the gray line is the forecast based on only recession periods in the full sample. Note that the recession sample forecasts use the recession-period coefficients to forecast through the end of 2009, even though the recession likely ended earlier.

Across all the measures in figure 5, the forecasts based on the full sample of data consistently underpredict the deterioration in labor market conditions. For example, the unemployment rate forecasted (panel A) at the end of 2009 lies roughly 2 percentage points below the actual unemployment rate, a finding noted by many commentators who worry that Okun's law no

FIGURE 5

Out-of-sample forecasts of labor market variables, 2007–09



Notes: Each panel is based on a simple statistical model that uses gross domestic product growth to generate out-of-sample forecasts of the labor market measure. All data from the first quarter of 1978 through the second quarter of 2007 are used to estimate the full sample forecasts, while data from just the recessions in this period are used to estimate the recession sample forecasts. The U-6 rate is a broader unemployment rate from the U.S. Bureau of Labor Statistics (see note 6 for details).
Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files; and Haver Analytics.

longer applies and labor markets are not functioning as in the past. However, if we use the recession sample (gray line), this simple activity model does a remarkably good job at forecasting the cumulative rise in the standard (panel A) and broader (panel B) unemployment rates and the fall in total payroll employment (panel C). That is, labor markets have mostly evolved about as we would expect given the severity of the recession.

But such a conclusion is not warranted for unemployment duration (panel D of figure 5).⁸ Forecasts based on both the full and recession samples fail to predict by up to over a month the dramatic rise in that series, starting in the fourth quarter of 2008. The remainder of this article is therefore focused on explaining the causes of the strikingly unusual increase in the length of unemployment spells.

Who are the long-term unemployed and how have they changed over time?

Figure 3 (p. 31) highlights the spike in average unemployment duration and long-term unemployment in 2009. It also illustrates that unemployment duration was already historically high going into the recent recession given the unemployment rate at the time. Relative to the black regression line that predicts duration based on the contemporaneous unemployment rate (figure 3, panel A), the black boxes there suggest that unemployment spells were already about four to five weeks higher, on average, than those during 1948–2007. For that reason, at least part of the explanation for current lengths of unemployment happened years ago.

Accordingly, table 1 examines the background characteristics of the long-term unemployed, in particular

gender, age, marital status, race, education, industry, and occupational background in 2009, in 1983 (when unemployment rates last reached 10 percent—and for the sake of comparison in the aftermath of a similarly severe recession), and in 2005–07 (before the start of the recent downturn). We also compare the distributions of these characteristics to their distributions in the entire labor force in the second set of columns.⁹ In the third set of columns, we report the ratio of the share of the long-term unemployed to the share in the labor force for each group. A number above 1 would imply that long-term unemployment was unconditionally more common in that group than would be expected given their representation in the labor force.

In the early 1980s, long spells of unemployment tended to be concentrated among factory and machine workers, who made up 29 percent of the labor force but 55 percent of the long-term unemployed, or nearly twice their representation in the work force (final row of table 1). Consequently, the long-term unemployed also tended to be heavily male (first column, first row) and only one in five long spells were from individuals with at least some college education (first column, fifteenth and sixteenth rows).

In 2009, factory and machine workers (and construction and manufacturing workers in general), males, and those with no college education still represented a larger share of the long-term unemployed than they did of the labor force (third column versus sixth column).¹⁰ However, the long-term unemployed became sectorally more diverse.¹¹ For example, in 2009, the long-term unemployed were more likely to come from professional and business services and finance, insurance, and real estate relative to 1983, while the share of manufacturing/factory workers went down. Generally, in 2009, long-term unemployment was more equally weighted across industry, occupation, education, gender, and age groups, and was therefore more representative of the labor force and the population than it had been two and a half decades ago.

Many important demographic shifts in the labor force have occurred concurrently with changes in the average length of unemployment. This has led several researchers (for example, Abraham and Shimer, 2002; Valletta, 2005; and Mukoyama and Şahin, 2009) to suggest a link between work force trends and unemployment duration. These links can be caused by differences in the propensity to be rehired in a timely fashion after job loss for particular demographic groups. For example, increases in college experience, as well as the general skills that education provides, might enable workers to be more adaptable and thus find job matches more quickly (of course, more job-specific

or industry-specific skills could potentially slow the process down).

In table 2, we provide a simple breakdown of changes in average duration of unemployment, using an approach called a Blinder/Oaxaca decomposition.¹² This decomposition enables us to estimate how much of the rise in unemployment duration is due to compositional changes in the pool of unemployed workers (for example, age, gender, education, and industrial composition); how much is due to longer spell lengths within each group (for example, longer spells among women or construction workers), holding the composition constant; and how much is due to interactions between changes in compositional effects and coefficients. We calculate these changes over two time periods roughly 20 to 25 years apart. First, we compare 1985–86 to 2005–06, when the economy was in the midst of expansions. Second, we examine two periods in our sample where unemployment was 10 percent or higher—the first six months of 1983 and the last six months of 2009 (that is, 1983:Q1–Q2 and 2009:Q3–Q4).

We find that most changes in the composition of the work force account for little of the increase in average duration of unemployment.¹³ The notable exception is the age structure of the population. Younger workers in the midst of a long unemployment spell tend to have shorter spells of unemployment than older workers in the same situation (Abraham and Shimer, 2002). Therefore, as the labor force has become older, average spells have tended to become longer. In table 2 (first column, second row), we show that changes in age can account for 0.7 weeks of the 1.3 increase in weeks from the mid-1980s to the mid-2000s, or about 53 percent. Yet, the changing age composition only accounts for about 25 percent of the rise in duration across the two periods of high unemployment (second column, second row). This suggests that as the baby boom generation continues to transition out of the labor force over the next decade, we should expect the average duration of unemployment to slowly fall.

The results of the decomposition also suggest that rising length of unemployment among women (holding the share of women in the labor force fixed) can account for virtually the entire increase in the average duration from the mid-1980s to the mid-2000s (table 2, first column, ninth row). This corresponds to the greater labor force attachment of women in recent decades and confirms Abraham and Shimer (2002), whose findings have a similar pattern. Change in unemployment duration within industries (first column, twelfth row) can also account for some of the secular pattern across expansions.

However, both the female and industry effects can explain a notably smaller share of the total

TABLE 1

Descriptive statistics: Long-term unemployed and labor force

		Long-term unemployed			Labor force			Ratio of long-term unemployed share to labor force share		
		1983	2005–07	2009	1983	2005–07	2009	1983	2005–07	2009
Gender	Male	67.6	55.5	58.6	56.0	52.4	52.2	1.21	1.06	1.12
	Female	32.4	44.5	41.4	44.0	47.6	47.8	0.73	0.93	0.87
Age	16–24	23.6	21.0	18.3	21.1	14.1	13.1	1.12	1.49	1.40
	25–54	65.5	62.1	63.4	64.9	67.7	66.4	1.01	0.92	0.95
	55–64	9.8	13.5	14.3	11.1	14.1	15.7	0.88	0.95	0.91
	65 and over	1.1	3.5	4.0	3.0	4.1	4.8	0.38	0.84	0.85
Marital status	Not married	50.9	65.7	62.4	39.0	43.8	44.1	1.30	1.50	1.42
	Married	49.1	34.3	37.6	61.0	56.2	55.9	0.81	0.61	0.67
Race	White	70.6	56.9	59.8	83.1	73.9	72.9	0.85	0.77	0.82
	Black	20.6	23.0	18.4	8.8	8.6	8.8	2.34	2.68	2.10
	Hispanic	3.1	8.5	6.9	3.1	6.5	6.7	0.99	1.30	1.02
	Other	5.6	11.6	14.9	5.0	11.0	11.6	1.14	1.06	1.29
Education	Less than high school	32.8	23.2	18.9	29.3	17.2	15.9	1.12	1.35	1.19
	High school graduate	46.3	35.6	37.1	37.9	30.9	30.3	1.22	1.15	1.23
	Some college	13.4	24.5	27.4	17.0	26.6	27.2	0.79	0.92	1.01
	College graduate	7.6	16.7	16.5	15.9	25.3	26.7	0.48	0.66	0.62
Industry	Agriculture, fishing, forestry, and mining	5.1	1.7	2.0	5.2	2.6	2.6	0.99	0.66	0.75
	Utilities and sanitation	0.8	0.6	0.7	1.4	1.1	1.2	0.62	0.55	0.58
	Construction	11.4	10.2	14.1	6.7	8.1	7.5	1.70	1.25	1.89
	Manufacturing	34.6	13.8	15.7	19.1	10.9	10.2	1.81	1.26	1.53
	Wholesale trade	3.2	2.6	2.3	4.1	3.0	2.6	0.78	0.88	0.90
	Retail trade	9.7	13.8	12.2	11.8	11.7	11.4	0.82	1.18	1.07
	Transportation and warehousing	4.5	3.8	3.9	4.0	4.2	4.1	1.13	0.90	0.94
	Finance, insurance, and real estate and leasing/rental	2.6	4.7	5.8	6.0	7.0	6.7	0.44	0.67	0.87
	Professional/business services and information	5.2	15.8	15.3	7.5	12.2	12.6	0.69	1.29	1.22
	Health, education, and social services	8.1	13.4	11.4	16.5	21.4	22.7	0.49	0.62	0.50
	Other services	11.1	16.3	14.4	12.9	13.0	13.5	0.86	1.26	1.07
	Government	3.0	3.0	1.8	4.8	4.8	4.9	0.63	0.62	0.35
Occupation	Executive and managerial	4.8	8.1	10.1	10.2	14.7	15.3	0.47	0.55	0.66
	Professional and technical	5.3	11.3	10.5	14.8	20.3	21.3	0.35	0.56	0.49
	Sales	6.3	11.7	11.5	11.3	11.4	11.0	0.56	1.03	1.04
	Administrative support	10.7	14.1	13.6	15.6	13.6	13.0	0.68	1.04	1.05
	Services	14.6	21.5	18.0	14.4	16.6	17.4	1.02	1.29	1.03
	Farming, forestry, and fishing	2.6	1.2	1.0	4.4	0.8	0.8	0.60	1.57	1.21
	Factory/machine workers	55.0	31.6	35.0	29.1	22.6	21.1	1.89	1.40	1.66

Notes: All values are in percent. Some columns may not total because of rounding.

Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files.

change in spell length when we compare the changes across the two periods of high unemployment in the second column of table 2. For example, changing coefficients for women can only account for about 35 percent of the rise in the average duration of unemployment across the two periods of high unemployment (second column, ninth row). Industry effects almost completely disappear (second column, twelfth row). Just under one-half (2.8 weeks out of 6.2 weeks) of the increase in duration from the first half of 1983 to the second half of 2009 is explained by direct shifts in composition and coefficients (second column, seventh and thirteenth rows).

Overall, the decomposition suggests that although demographic factors can account for much of the secular increase in unemployment duration, they can only account for a portion of the especially sharp rise in durations that has accompanied this most recent recession. This suggests that other factors must be driving this phenomenon—the topic that we turn to next.

Labor market transitions, the unemployment rate, and unemployment duration

In order to better understand the causes of the recent sharp rise in long-term unemployment, it is useful to develop a framework for studying labor market dynamics during the business cycle. In this section, we begin formulating this framework by showing how movements between being employed, unemployed, and out of the labor force (labor market transitions) have contributed to cyclical patterns in unemployment historically and during the most recent recession. We then generate a model that uses labor market transitions to create counterfactual scenarios that would correspond to alternative views of what may be driving labor markets. Finally, we use this apparatus to provide some insight into the causes of long-term unemployment. We also use these results later to analyze the implications of long-term unemployment for the aggregate economy going forward.

To measure labor market transitions, we exploit the fact that the CPS interviews whatever household unit is living at a particular address for four consecutive months, skips the address for eight months, and then returns for more interviews over four consecutive months. This allows us to track many household units

over time. We follow previous studies that have used matching algorithms to identify individuals who are living at the same address in consecutive months and build a panel data set containing the labor market status of individuals at multiple points in time. Specifically, we consider the nine possible transition probabilities (transition rates) across the states of employment (*E*), unemployment (*U*), and out of the labor force (*O*).

Transition rates and the unemployment rate

Figure 6 plots these nine seasonally adjusted monthly transition rates (blue lines), along with six-month moving averages of each to smooth out some of the noise in the data (black lines). Two key transitions for explaining past changes in the unemployment rate are movements from employment to unemployment (*EU*) and unemployment to employment (*UE*).¹⁴ The *EU* transition rate measures the fraction of employed individuals who separate from their employer and move into unemployment. We will hereafter refer to this as the “separation rate.”¹⁵ The *UE* rate is sometimes referred to as the hiring rate. Shimer (2007) has argued that most of the variation in the unemployment rate is due to fluctuations in the hiring rate rather than the separation rate, although this conclusion has been disputed by Elsby, Michaels, and Solon (2009), who argue that both rates have been of significant importance.¹⁶

TABLE 2

Decomposition of the secular change in the average duration of unemployment, 1980s to 2000s

	1985–86 to 2005–06	1983:Q1–Q2 to 2009:Q3–Q4
Total change to explain	1.3	6.2
Due to changes in composition		
Age	0.7	1.6
Gender	–0.1	0.0
Race	0.0	0.1
Education	–0.1	–0.5
Industry	0.0	0.0
Total	0.5	1.2
Due to changes in coefficients		
Age	0.0	–0.3
Gender	1.6	2.1
Race	–1.0	0.0
Education	–0.3	–0.3
Industry	1.3	0.1
Total	1.6	1.6
Interactions between changes in composition and coefficients	–0.8	3.3

Notes: All values are in weeks. See note 12 for further details. The second column does not total because of rounding.

Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files.

Movements in the *EU* have been particularly pronounced in this recession: The separation rate has risen by nearly 70 percent. This disproportionate hike is shown more clearly in figure 7 (p. 40), where we compare the proportional change in the *EU* and *UE* transition rates in the current business cycle with the recessionary periods in 1981–82 and 2001. The *EU* transition rate followed its historical pattern during 2008 but then began rising sharply early in 2009. Relative to the acceleration in the *EU* rate, the *UE* rate appears to have fallen more gradually, though proportionately more than in previous recessions.

To assess how important the transitions out of employment versus transitions out of unemployment have been in explaining the rise in the unemployment rate during the most recent downturn, we perform some simple simulations. We start with the actual levels of those who are employed, unemployed, and out of the labor force and the smoothed values of all nine of the labor market transition rates at the end of 2007. We then use the actual transition rates starting in January 2008 to simulate the new counts of individuals in each labor market state for each month going forward. This is described in greater detail in box 1 (p. 41). With some basic adjustments, we are able to match the actual monthly unemployment rates through the end of 2009 almost perfectly.

We then conduct the following two experiments. First, we hold all transition rates constant at their December 2007 values except for the three transitions that start with the employment state in the initial month (*EE*, *EU*, and *EO*).¹⁷ Those transition rates are allowed to vary according to what actually transpired in 2008 and 2009. In essence, this exercise, which is plotted as the dark blue line on figure 8 (p. 42), captures the effects of transitions out of employment into non-employment (being either unemployed or out of the labor force) on the aggregate unemployment rate.¹⁸ Analogously, we do a second experiment where only the transitions from the state of unemployment (*UE*, *UU*, and *UO*) are allowed to change. This captures the effects of the fall in the exit rate out of unemployment into being either employed or out of the labor force. Those results are shown as the light blue line in figure 8. The black line is the actual unemployment rate, and the gray one is the actual unemployment rate in December 2007.

We find that the changes in the transition rates out of employment (all else being equal) would only raise the unemployment rate by 1 percentage point by the end of 2009. In contrast, changes in the transition rates out of unemployment would raise the unemployment rate by 2.2 percentage points. Broadly speaking,

this suggests that the combined effects of moving out of unemployment (*UE*, *UU*, and *UO*)—including, prominently, the transition into a job—explain more of the actual increase in the unemployment rate over the past two years than the combined effects of moving out of employment (*EE*, *EU*, and *EO*).¹⁹

Transition rates and unemployment duration

We next turn to using these exercises to explain unemployment duration. The simulation is similar as before except that we now explicitly incorporate the distribution of unemployment duration into the analysis by using five-week “bins” of unemployment spells (that is, 0–4 weeks, 5–9 weeks, and so on). We start with the distribution of unemployment duration at the end of 2007 and use estimates of the actual transition rates into and out of unemployment for each bin, along with estimates for the other transition rates, to update the distribution of duration each month. We again find that the simulation does extremely well at replicating the sharp rise in the average duration of unemployment during 2009.²⁰

In figure 9 (p. 42), we show that if only the *EE*, *EU*, and *EO* followed their actual paths and all the other transition rates stayed constant at their December 2007 values, the average duration of unemployment would have only increased slightly, to about 19 weeks by the end of 2009 (dark blue line). If, however, only *UE*, *UU*, and *UO* followed their actual paths and the other transition rates stayed flat, unemployment duration would have increased to nearly 23 weeks (light blue line). So it appears that for both the unemployment rate and the average duration of unemployment, transition rates from the starting state of unemployment have been the important driving influences.²¹

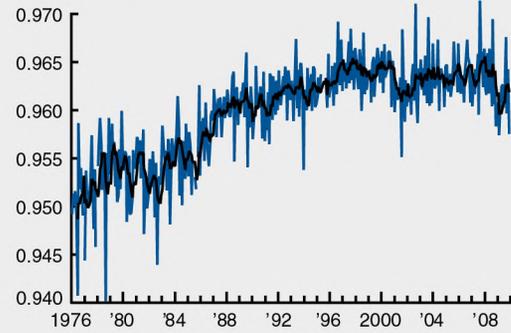
Simulated effects of federal unemployment insurance benefit extensions

As noted previously, the spike in the average duration of unemployment starting in mid-2009 is hard to explain using demographics or the standard association with deteriorating GDP growth. One plausible explanation is the unprecedented extension of unemployment insurance benefits. The maximum number of weeks of eligibility rose from 26 weeks to 39 weeks in July 2008 with the passage and creation of the Emergency Unemployment Compensation (EUC) federal program. Since then, extensions have risen at varying rates, depending on the unemployment situation of individual states.²² Figure 10 (p. 43) plots the weighted national average of the maximum number of weeks of unemployment benefit receipt allowed (blue line); the weights for this average are based on the size of the unemployment pool in each state. As of January 2010,

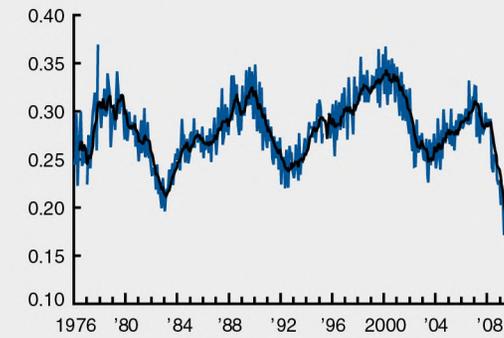
FIGURE 6

Labor market transition rates, 1976–2009

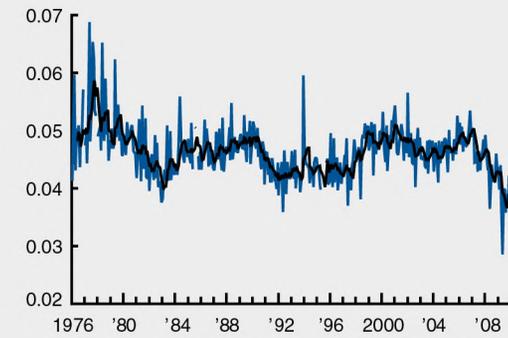
A. Employment to employment transition rate



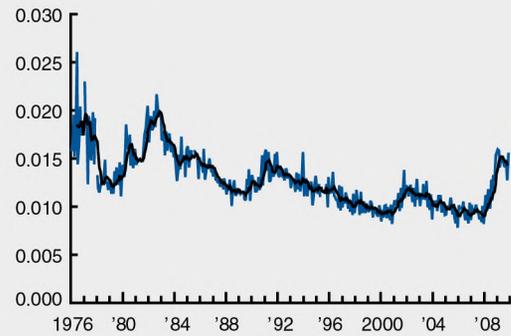
B. Unemployment to employment transition rate



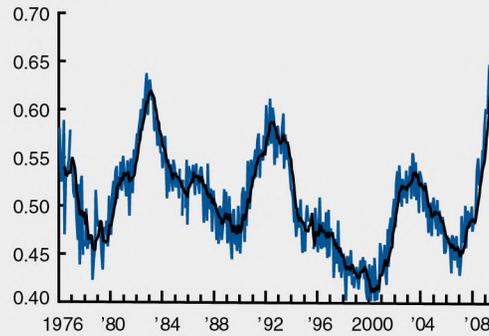
C. Out of the labor force to employment transition rate



D. Employment to unemployment transition rate



E. Unemployment to unemployment transition rate



F. Out of the labor force to unemployment transition rate

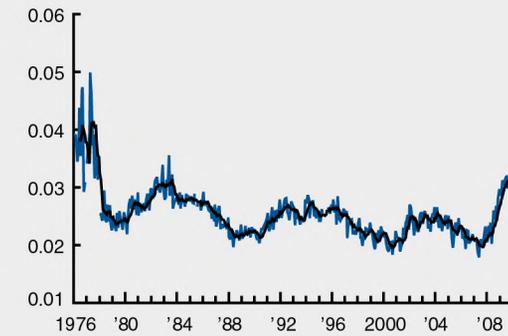
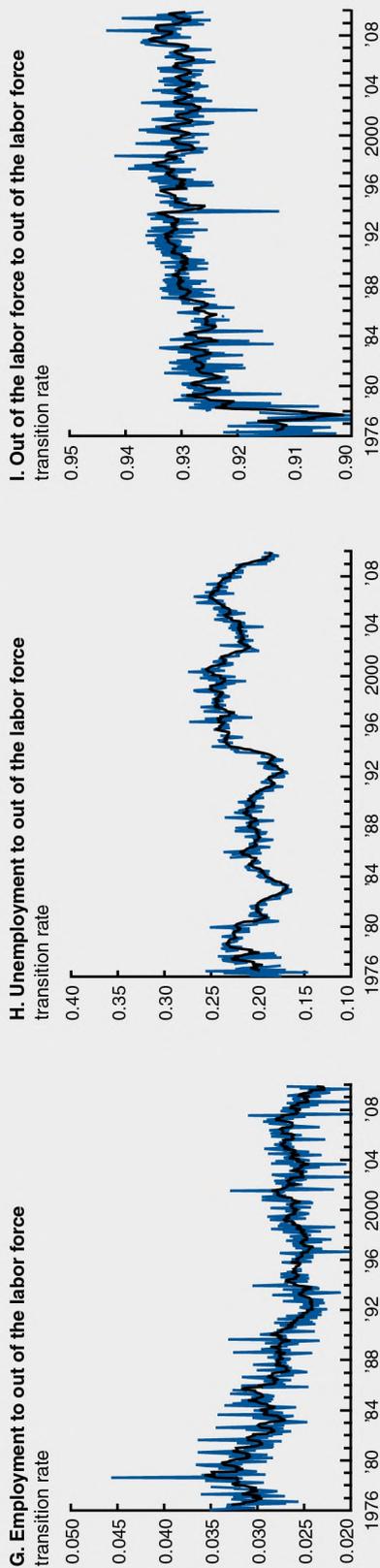


FIGURE 6 (CONTINUED)

Labor market transition rates, 1976–2009



Notes: All transition rates are seasonally adjusted and plotted as blue lines, and their six-month moving averages are plotted as black lines. Transition rates are not available for some months, when matching individuals was not possible because of survey redesigns.

Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files.

unemployed workers in 14 states were allowed the maximum of 99 weeks of UI benefits and the national average was 90 weeks. By contrast, in 1983, the maximum potential duration of UI coverage in any state had reached 55 weeks.

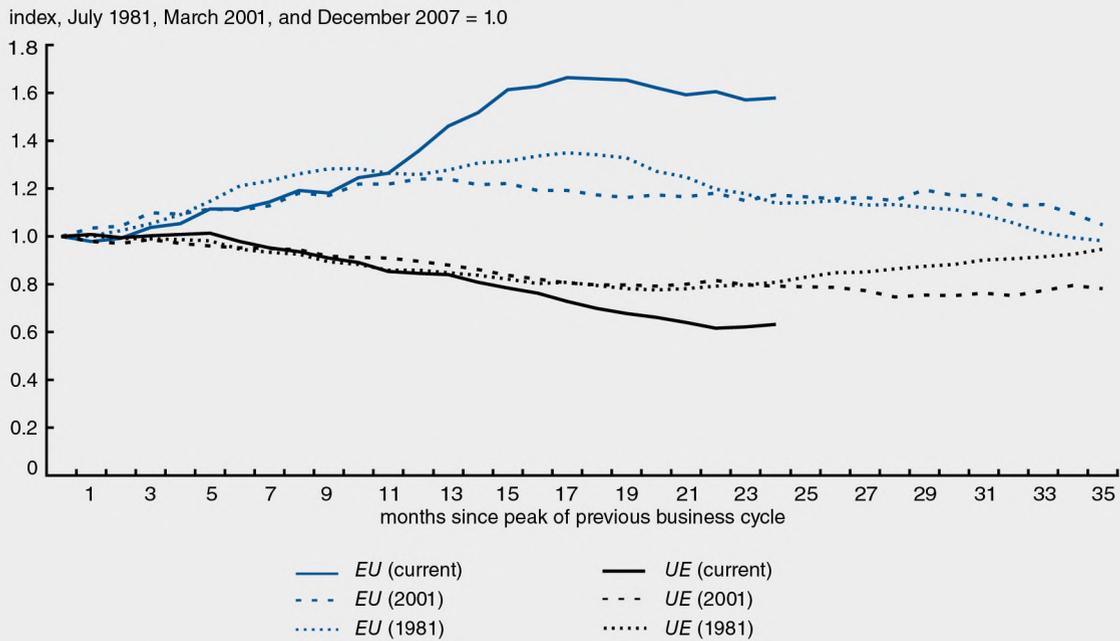
In order to estimate the possible effect of UI benefit extensions on unemployment duration, we use previous studies of the effect of an additional week of maximum benefits on average duration. A prominent example in this literature is Katz and Meyer (1990), who use a rich statistical model and administrative data from the UI system to estimate the probability of leaving unemployment during the early 1980s recessions. They identify the impact of UI through variation in maximum benefits both within and across states that shift as a result of eligibility rules and legislative changes. They find that the average duration of unemployment rises by 0.16 weeks to 0.2 weeks for each additional week of benefits extended.

Katz and Meyer (1990) face the difficult problem of disentangling the effects of UI benefit extensions from the effects of poor economic conditions that typically prompt benefit extensions in the first place. When the economy is in a recession, longer spells of unemployment are expected irrespective of the generosity of the unemployment insurance program. To get around this problem, Card and Levine (2000) use an increase in the maximum number of weeks of benefit eligibility in New Jersey in 1996; this increase was unrelated to the state of the economy at the time. In fact, this particular extension, which was driven by political considerations, took place in the midst of an expansion and therefore might be less susceptible to the bias of recession-driven extensions. Indeed, they find a smaller effect than Katz and Meyer (1990) and much of the rest of the literature; mean duration rises by about 0.1 weeks for each additional week of benefits. In order to reflect our uncertainty over the true effect, we use both estimates.

We begin our analysis in June 2008, when maximum UI eligibility was 26 weeks and unemployment spells lasted about 17 weeks on average (the six-month mean from January through June 2008). We then calculate an estimated effect of the extension in unemployment benefits for each subsequent month beginning with July 2008.²³ For such a calculation, two additional inputs are required. First, we need the share of the unemployed who are actually receiving benefits because they are the only ones who would be directly affected by policy changes. The black line in figure 10 (p. 43) shows that the

FIGURE 7

Employment-to-unemployment (EU) and unemployment-to-employment (UE) transitions over selected business cycles



Notes: The most recent business cycle peak was in December 2007, according to the National Bureau of Economic Research; July 1981 and March 2001 were the business cycle peak months corresponding with the recessions beginning in 1981 and 2001, respectively. These three dates correspond with month 0 in this figure.
 Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files.

share of the unemployed receiving UI surged from 41 percent in July 2008 to 67 percent in December 2009. Second, we must assume a time period over which to distribute the full effect of the extension in benefits. We distribute the effects of the initial 13-week extension of benefits that took place in July 2008 over a full year.²⁴ The additional extensions that increased the maximum potential duration in the UI system beyond 52 weeks, beginning in December 2008, are spread over two years because the much larger extensions are likely to alter behavior over a longer period of time.

Using these inputs, our estimates based on the Katz and Meyer (1990) elasticity suggest that the extension of UI benefits during 2008 and 2009 may account for as much as 3.1 weeks of the 12-week increase in the average duration of unemployment that took place over this period. The estimate based on the Card and Levine (2000) analysis suggests that it could explain about 1.6 weeks. These assessments, which we consider our range of preferred estimates, suggest that the effect of unemployment insurance extensions on the average duration of unemployment is on the order of 10–25 percent of the total increase since July 2008.

Alternatively, if we spread the effect of the extensions beginning in December 2008 over just one year, this would raise our estimates of the contributions to between 3 weeks and 6 weeks, or between 25 percent and 50 percent. It is also important to note that our calculations have not considered the potential effect of the “reach back” provision in the EUC program that allowed extensions for those who had exhausted their unemployment benefits as early as May 1, 2007. It is possible that this provision could further raise our estimates of the impact of UI benefits.

Effects of unemployment insurance benefit extensions on the unemployment rate

We can also utilize the transition data to examine whether movements from unemployment to out of the labor force (*UO*) and those in the opposite direction (*OU*) yield additional clues about possible changes in classification between the non-employed that may have arisen as a result of UI benefit extensions. In figure 11, panel A (p. 44), we plot all the possible transitions from unemployment during the current business cycle (blue lines) and compare them with

BOX 1

Methodology for simulating paths of the unemployment rate and unemployment durations

In this article, we use transitions across different labor market states as a tool for simulating the paths of the unemployment rate and unemployment durations. This allows us to consider alternative scenarios for the path of the unemployment rate or durations either historically or going forward based on changing the paths of particular transitions. Using this approach, we can infer the relative importance of particular economic phenomena that are related to certain transitions as described in the text. An important caveat is that this is a mechanical approach that may or may not correspond well to changes in the actual economy. For example, conditions that may change a particular transition rate may also affect other transition rates in ways that we may not consider.

If we consider time as discrete and denote it as t and let x stand for a particular labor market state, that is, employed, unemployed, or out of the labor force, ($x = \{E, U, O\}$), then each of the nine possible transitions are defined as the probability of being in a particular labor market state in t conditional on having been in the same or different labor market state in the prior period. For example, the EU transition is defined as:

$$EU = \text{Prob}(x_t = U | x_{t-1} = E).$$

For each initial state in $t - 1$, the three possible transitions must sum to 1. For example, $EE + EU + EO = 1$. Each of the nine transitions for each month is estimated empirically using the matched *Current Population Survey* as described in the text. To implement the simulation we start by inputting the levels of those who are employed, unemployed, and out of the labor force for a chosen base period. We then simulate the next period's level of E , U , and O , using the assumed levels for the base period and an assumed path for the transition probabilities. For example, if we wish to simulate the actual path of the unemployment rate through 2008 and 2009, we define December 2007 as our base period and then use the actual estimated values of the transition probabilities for January 2008 through December 2009.¹

For example:

$$E_{\text{Jan08}} = E_{\text{Dec07}} \times EE_{\text{Jan08}} + U_{\text{Dec07}} \times EU_{\text{Jan08}} + O_{\text{Dec07}} \times EO_{\text{Jan08}}.$$

We use several methods to pose alternative transition rates, depending on our question of interest. To address the relative importance of transitions from employment versus transitions from unemployment, we start with a baseline path where all of the transition rates are constant. We then change the paths of all three transition rates from either employment or unemployment simultaneously. For example, we simulate the effects arising only from changes from the employment state by changing the paths of EE , EU , and EO simultaneously.

A second approach is used when we wish to hold the UO and OU transition rates fixed at a particular rate. In this case, we allow the UE and OE rates to follow their actual paths and then adjust the UU and OO so that the probabilities from U and from O each sum to 1. Finally, for the simulation that attempts to reproduce the forecast of the unemployment rate according to the *Blue Chip Economic Indicators*, we assume that the EU , EE , UU , and UE rates take five years to return to their historical average values. We then adjust the EO and UO rates so that the three transitions from E and from U sum to 1.

¹Rather than immediately going from the base period to the first period of the simulation, we first use the transition rates from the base period and run about ten iterations of the model so that the values of E , U , and O and the implied unemployment rate reach a steady state, where they are unchanging. We then proceed to use the steady-state values for the simulation. The steady-state values may differ from the actual values in the base period. For example, the steady-state value of the unemployment rate in December 2007 is about 80 percent of the actual value. We therefore scale the subsequent values of the simulation by a factor of 1.25. This discrepancy is likely due in part to the inability to account for month-to-month compositional changes that arise from the fact that individuals enter or exit the working age population. Measurement error and differences between the complete population and the matched sample may also play a role. This approach assumes that although we cannot match the level of unemployment, we can match changes over time.

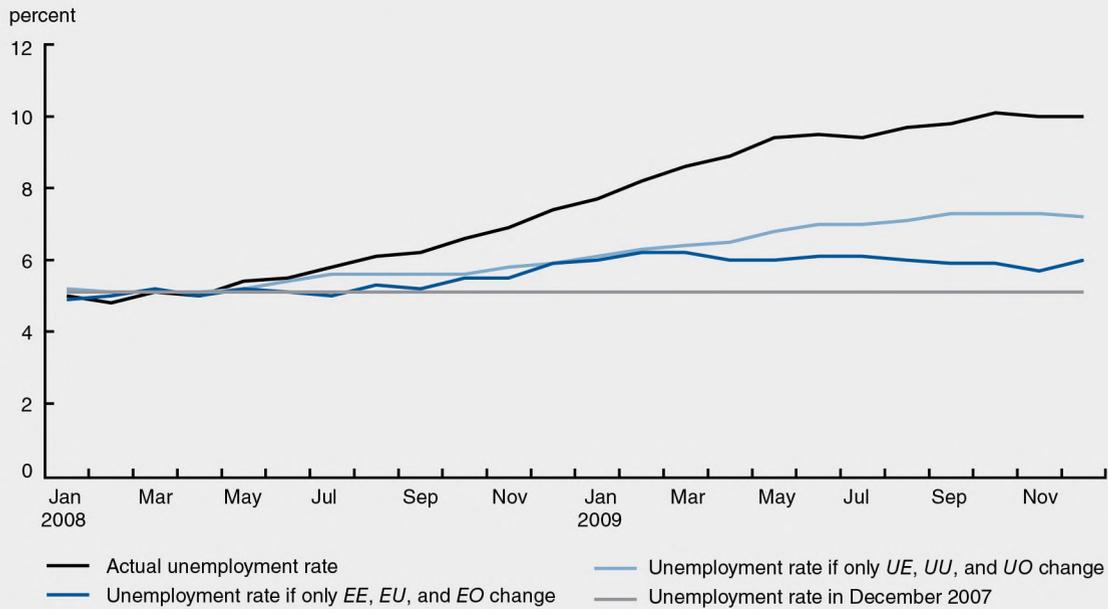
those during the 2001 recession and its aftermath (black lines). The series with the boxes represent the UE transition rates and are identical to those shown in figure 7 (p. 40). To this we add UU transition rates (diamonds) and UO (circles) rates. It appears that the UU and UO rates in the current recession track the rates in 2001 reasonably well for the first 16 or so months of the downturn before beginning to diverge. In contrast, the UE rates diverge earlier in the cycle. One

possible reason for this pattern is that individuals who would have normally dropped out of the labor force at this point in the cycle chose to remain unemployed—perhaps to continue to collect unemployment benefits.

In figure 11, panel B (p. 44), we focus only on the rate of UO transitions and add data from the 1981–82 recession (dotted). This panel shows that the UO path during the current recession resembles the UO path during the 1981–82 recession reasonably well, suggesting

FIGURE 8

Counterfactual effects of changing labor market transition rates on the unemployment rate, 2008–09

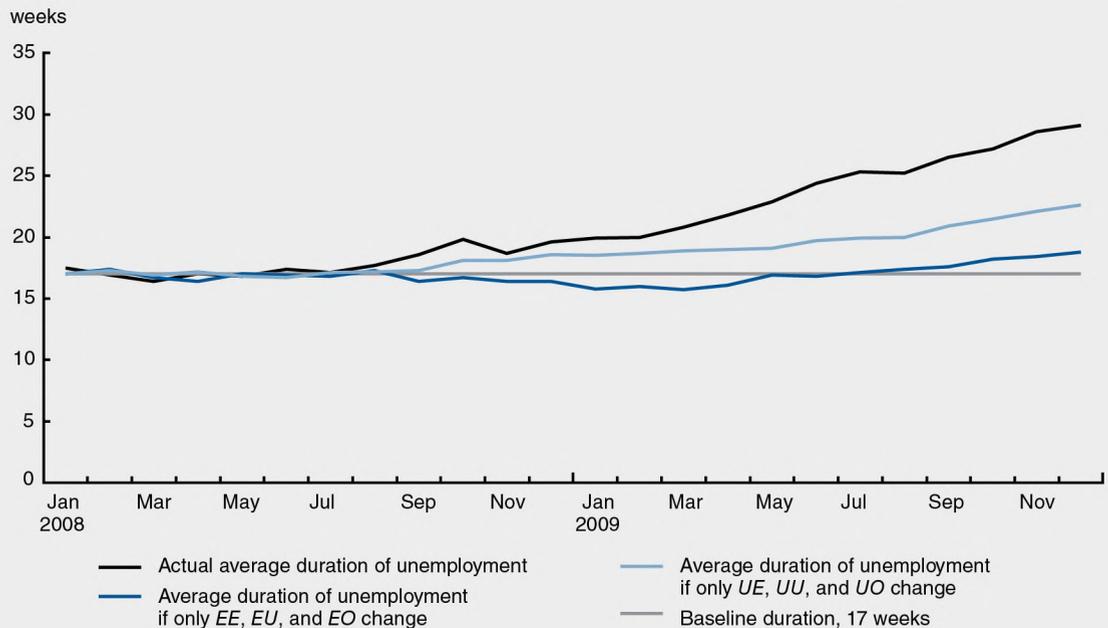


Note: *E* indicates employment, *U* indicates unemployment, and *O* indicates out of the labor force.

Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files; and Haver Analytics.

FIGURE 9

Counterfactual effects of changing labor market transition rates on the duration of unemployment, 2008–09

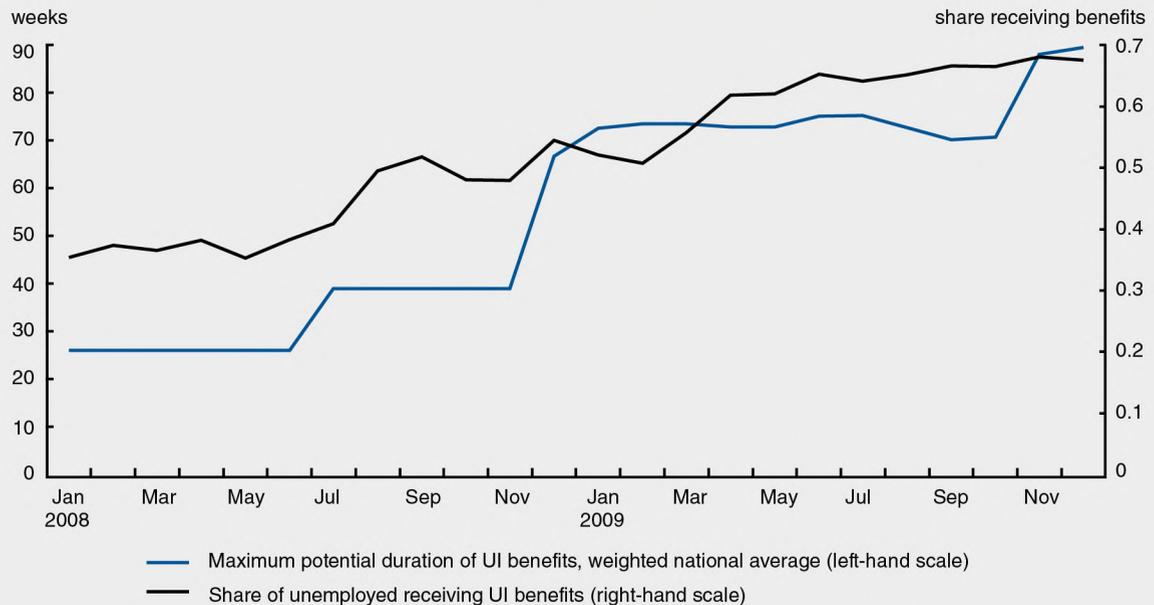


Notes: *E* indicates employment, *U* indicates unemployment, and *O* indicates out of the labor force. The average duration of unemployment in December 2007 was about 17 weeks, so we use this duration as our baseline.

Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files.

FIGURE 10

Maximum potential duration of unemployment insurance (UI) benefits and the share of unemployed receiving benefits, 2008–09



Notes: The weights for the national average duration are based on the size of the unemployment pool in each state. The share of unemployed receiving UI benefits is the number of individuals with continuing claims under federal and state unemployment insurance programs divided by the number of unemployed; this fraction is seasonally adjusted.
Sources: Authors' calculations based on data from the U.S. Department of Labor, the U.S. Bureau of Labor Statistics, and Haver Analytics.

that the departure from the 2001 pattern may simply reflect the greater severity of the current recession. That said, figure 11, panel C suggests that the rate of *OU* transitions in the current recession appears to move substantially higher in percentage terms than the patterns observed during the previous downturns.

We conduct a simulation motivated by figure 11 to ask how different the unemployment rate would be had the paths of the *UO* and *OU* transitions stayed constant at their values 16 months after the start of the recession (April 2009). In order to ensure that the probabilities from a particular state add up to 1, we allow the *UE* and *OE* rates to follow their actual paths and adjust the *UU* and *OO* rates so that the probabilities sum to 1. Figure 12 shows that under this counterfactual scenario the result of this exercise would be to lower the unemployment rate to 9.3 percent as of December 2009—about 0.7 percentage points below the actual unemployment rate that month. Although this is a relatively crude and mechanical approach, it nonetheless provides a magnitude for the possible effect of unemployment insurance benefit extensions on the unemployment rate.

Implications of long-term unemployment for the aggregate economy

In this section, we consider how the increase in long-term unemployment may affect the economy going forward. We consider the effects of the unemployment duration structure on the unemployment rate and then on compensation growth.

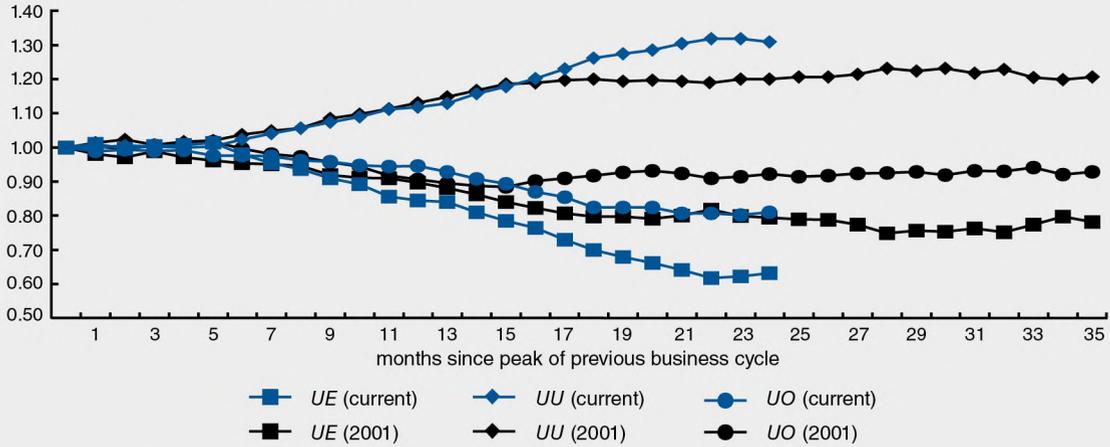
Effects of duration structure on the unemployment rate

As unemployment spells lengthen, the probability of finding a job in a given time period declines—an association that is robust across time and demographic groups. The pattern is illustrated in figure 13 (p. 46), which plots the probability of being employed today for various lengths of unemployment duration in the previous month (horizontal axis). For example, at 0–4 weeks of unemployment, the average probability of finding a job in the following month is 34 percent, but at 25–29 weeks, it is only 19 percent.²⁵ As much as this phenomenon is due to diminished job skills and weakened social networks, it could have a real impact on the labor market recovery while the broader economic recovery takes hold.

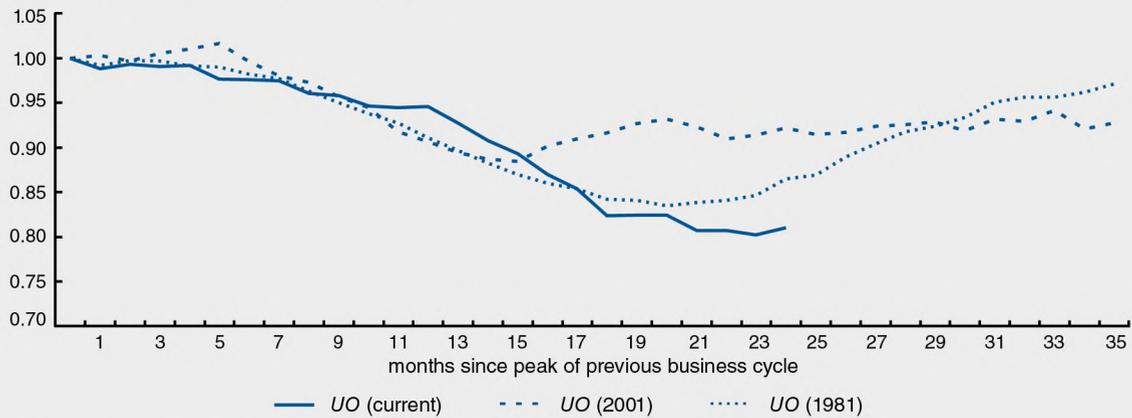
FIGURE 11

Labor market transitions from non-employment

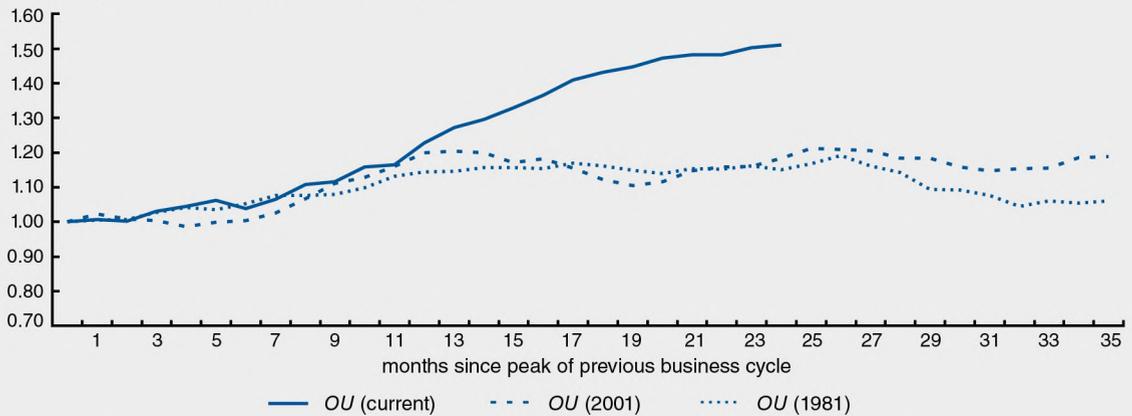
A. Transitions from unemployment, current recession versus 2001 recession
index, March 2001 and December 2007 = 1.0



B. Transitions from unemployment to out of the labor force (UO)
index, July 1981, March 2001, and December 2007 = 1.0



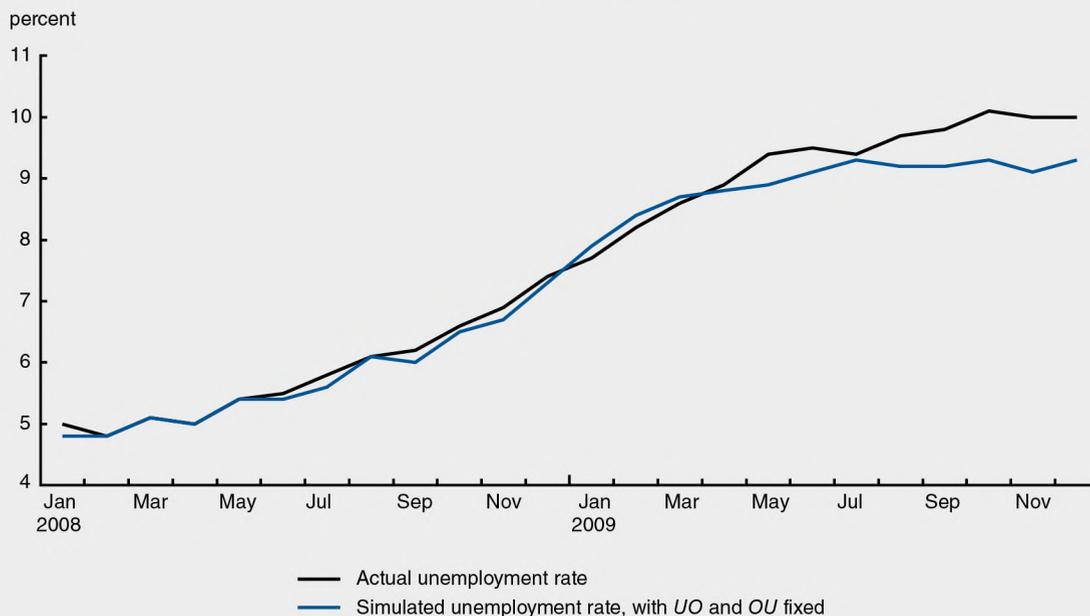
C. Transitions from out of the labor force to unemployment (OU)
index, July 1981, March 2001, and December 2007 = 1.0



Notes: *E* indicates employment, *U* indicates unemployment, and *O* indicates out of the labor force. The most recent business cycle peak was in December 2007, according to the National Bureau of Economic Research; July 1981 and March 2001 were the business cycle peak months corresponding with the recessions beginning in 1981 and 2001, respectively. These three dates correspond with month 0 in this figure. Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files.

FIGURE 12

Counterfactual effects of holding unemployment-to-out-of-the-labor-force (*UO*) and out-of-the-labor-force-to-unemployment (*OU*) transition rates fixed from April 2009 onward on the unemployment rate, 2008–09



Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files; and Haver Analytics.

In order to investigate this possibility, we use our transition rate model but substitute aggregate transition rate probabilities for movement from unemployment (*UE*, *UU*, and *UO*) with analogous transition rate probabilities for each five-week bin of unemployment duration. We start by simulating a baseline path that roughly matches the January 10, 2010, forecast of the unemployment rate through 2011 according to the *Blue Chip Economic Indicators* (Blue Chip), a survey of America's top business economists (Aspen Publishers, 2010). We then pose an alternative path where the only change is to make the share of the unemployed in each five-week bin at the beginning of the simulation (January 2010) match their mean historical values. In figure 14, panel A, we show that this alternative initial distribution of duration would immediately lower the unemployment rate by about 0.4 percentage points relative to the Blue Chip path. We find, however, that duration quickly reverts back to high levels (figure 14, panel B) and that the unemployment rate path converges to what it would have been had the model started with the actual distribution of duration. The main lesson we take from this exercise is that the unemployment rate is probably about half a point higher than it would be if unemployment spell lengths were at more historical levels.

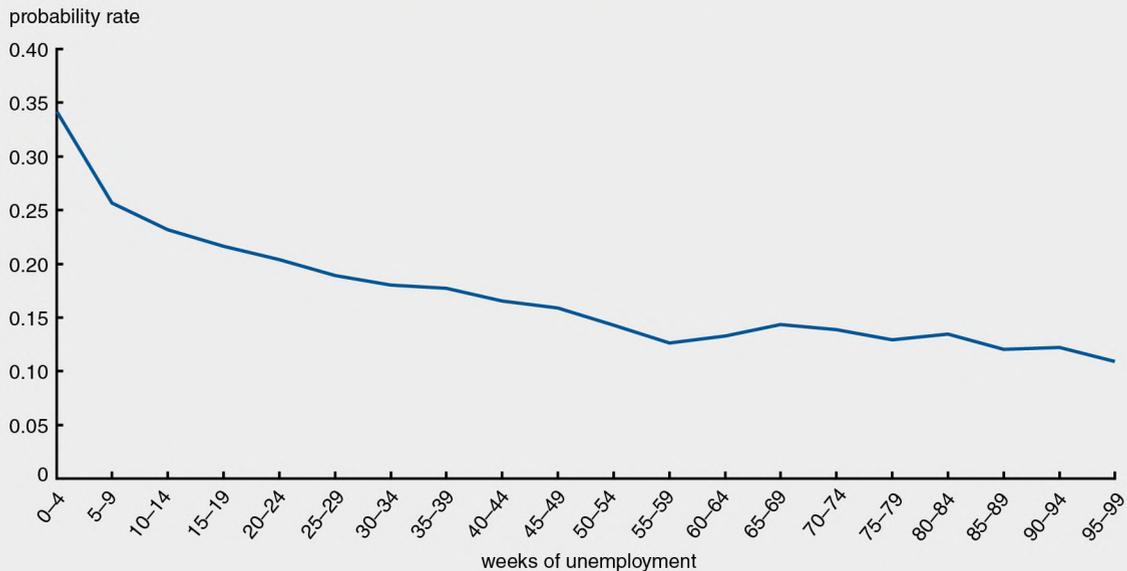
Effects of duration structure on compensation growth

Lastly, we consider the possible effects of higher long-term unemployment rates on aggregate wage growth. It is not obvious a priori what the expected effects should be. If the long-term unemployed are readily employable and can fulfill vacancies, then there is a sense in which they may be more eager to return to work at the prevailing wage than individuals with short unemployment durations. In this case, the long-term unemployed may reduce wage pressures. If, however, many of the long-term unemployed are more akin to individuals who have stopped searching for work and have left the labor force, perhaps because of a geographical or skills mismatch, then they may play little role in bidding down wages.

Since this is ultimately an empirical question, we undertake a simple exercise using Phillips curve style regressions to address this. We use data on year-over-year growth in real compensation per hour. Figure 15 (p. 48) shows that, as expected, there is a negative relationship between compensation growth and the unemployment rate. The black boxes signify the values starting with 2008:Q1, when the recession began. We regress compensation growth on the unemployment rate for the post-1975 period and calculate the predicted values.

FIGURE 13

Probability of reemployment, by unemployment duration



Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files.

We then add the share of the unemployed in each of the five-week bins of unemployment duration to the regression and reestimate the model. We plot both sets of the forecasted values, along with the actual growth rate of real compensation, in figure 16. We find that there is little difference in magnitude between the two forecasts. For much of the past 20 years, the predictions that incorporate unemployment duration are slightly lower than those that do not. However, there is little economically important difference in the most recent period. Overall, this suggests that at least for predicting aggregate compensation trends, there is no clear-cut indication that rising unemployment duration will signify any more or less slack than the information contained in the unemployment rate. This might be because rising unemployment duration produces countervailing forces on wage pressures as hypothesized earlier. However, the statistical model used to estimate these relationships is based on historical associations, whereas the current distribution of unemployment spell length is unprecedented.

Conclusion

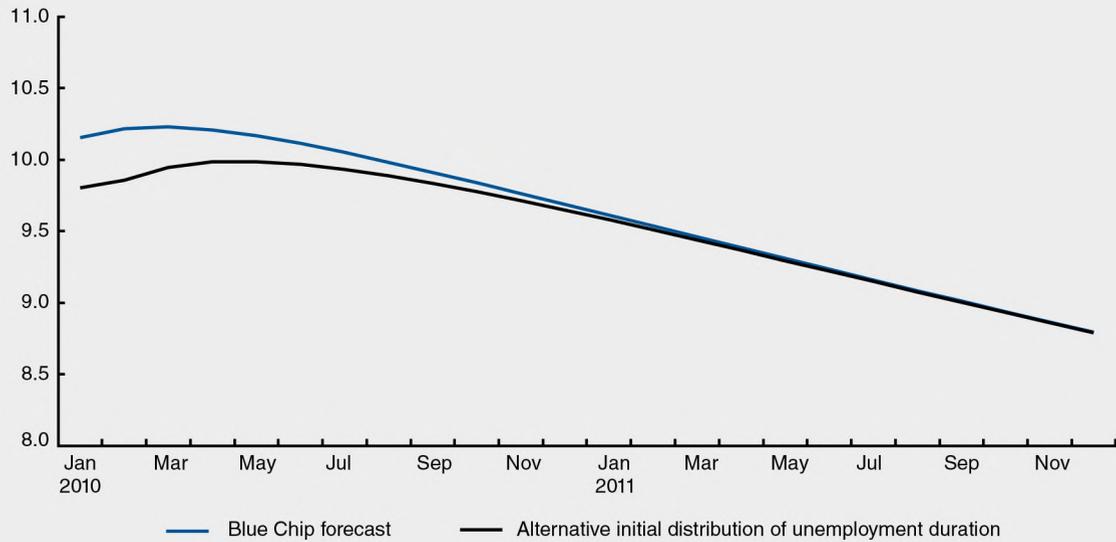
The average length of an ongoing spell of unemployment topped 30 weeks in December 2009, with more than 40 percent of the unemployed out of work for over six months. These numbers far exceed anything

recorded in the post-World War II era. In this article, we analyze the factors behind this historically unprecedented rise in long-term unemployment and explain what it might imply for the economy going forward. We show that roughly half of the rise relative to previous deep modern recessions was due to demographic factors in place well before the recession began. The remaining unexplained increase is due primarily to especially weak labor demand, reflected in low levels of hiring. Perhaps 10–25 percent of the increase in long-term unemployment from mid-2008 to the end of 2009 is associated with extensions of unemployment insurance benefits. These estimates for the current business cycle constitute a notable departure from historical patterns in transitions between being unemployed and out of the labor force. Some simple counterfactual estimates suggest that had these transitions followed more typical patterns, the unemployment rate might be about 0.7 percentage points lower. Finally, we find that high levels of long-term unemployment typically persist well into an economic recovery, since firms tend to hire the long-term unemployed last. Some simple simulations suggest that a historically long unemployment duration distribution as currently experienced in the United States could slow the process of labor market recovery, but it is not expected to have much of an impact on compensation growth.

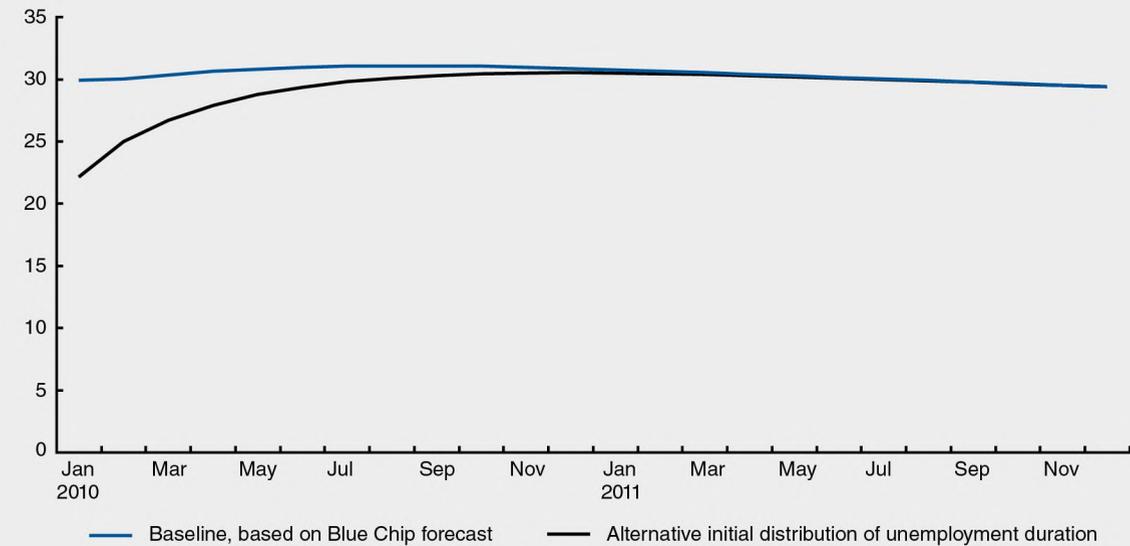
FIGURE 14

Simulated effects of changing the initial distribution of unemployment duration, 2010–11

A. Effects on the January 2010 Blue Chip forecast of the unemployment rate
percent



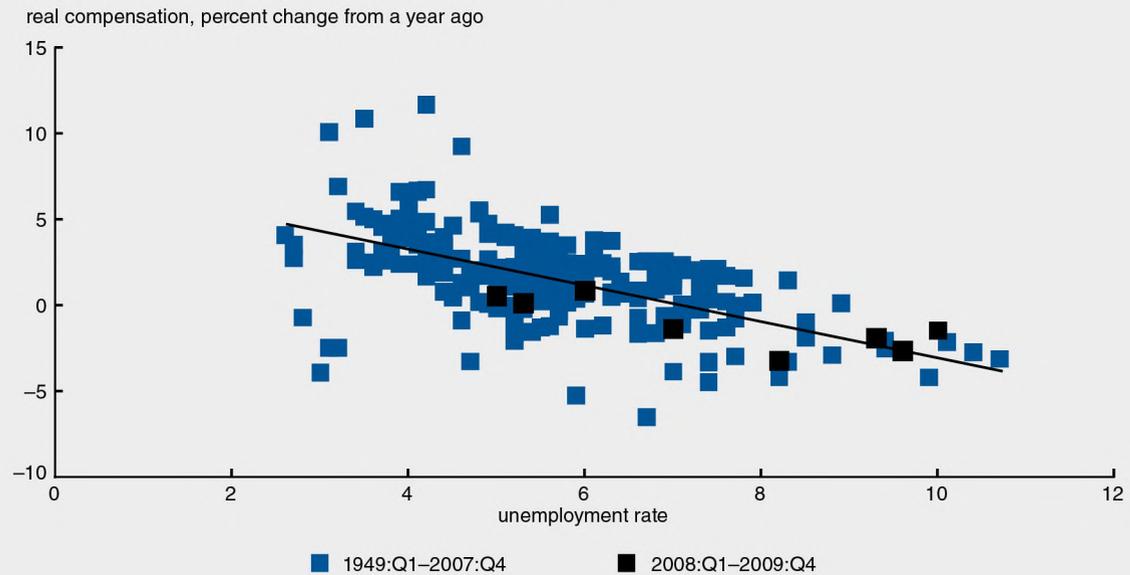
B. Effects on the path of the average duration of unemployment
weeks



Note: Blue Chip forecast refers to the forecast of the unemployment rate through 2011 according to the *Blue Chip Economic Indicators*.
Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics, *Current Population Survey*, basic monthly files; and Aspen Publishers (2010).

FIGURE 15

Real compensation growth versus the unemployment rate, 1949–2009

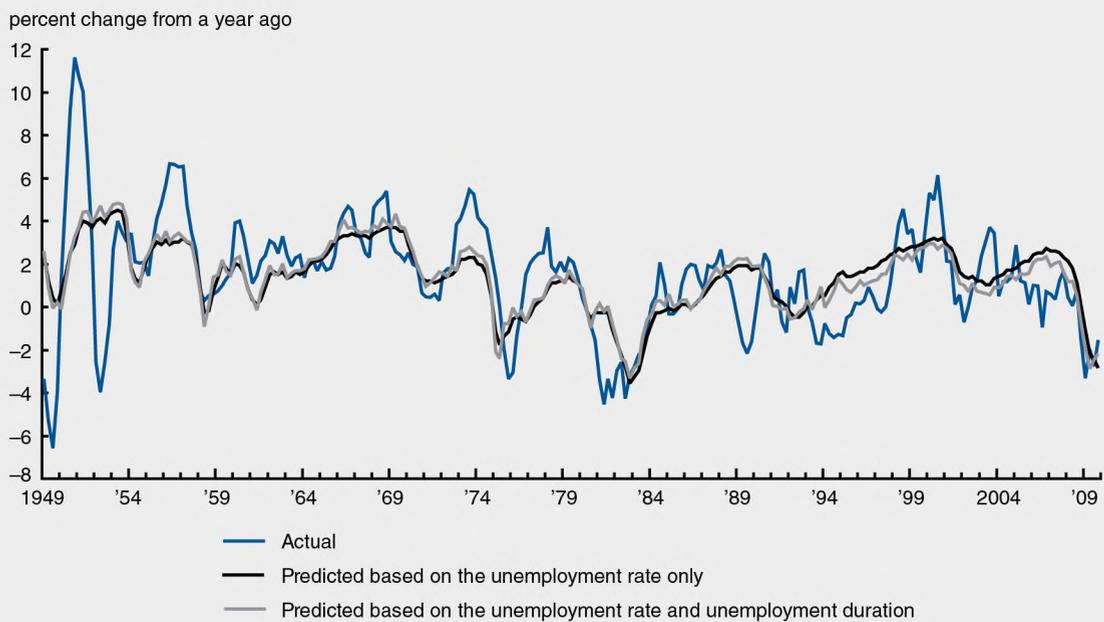


Note: The black line represents the relationship between the unemployment rate and the percent change of compensation from a year ago over the period 1949–2009.

Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

FIGURE 16

Real compensation growth, actual versus predicted, 1949–2009



Source: Authors' calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

NOTES

¹Alternatively, the relationship between the length of time out of work and the diminishment of work prospects could be picking up unobserved differences in worker quality between those who are unemployed for short and long spells (Ham and Rea, 1987; Kiefer, 1988; and Machin and Manning, 1999). In this case, longer spells in and of themselves do not lead to worse outcomes. It is very difficult to convincingly identify which of these channels dominates without strong statistical assumptions.

²Based on transition patterns between being employed, unemployed, and out of the labor force altogether, we estimate that UI extensions increased the unemployment rate by roughly 0.7 percentage points during 2008–09.

³Long-term unemployment is a good deal less common in the United States than in much of the developed world (for example, Machin and Manning, 1999). As of 2008, the last year for which comparable data are available, the share of the unemployed out of work more than six months was two times, and in some cases four times, higher in Belgium, the Czech Republic, France, Germany, Greece, Hungary, Italy, Luxembourg, the Netherlands, Portugal, Switzerland, the United Kingdom, and Japan.

⁴The most recent numbers from the *Current Population Survey* are still well below the prevalence of long-term unemployment during the Great Depression. Unfortunately, national data on unemployment duration before World War II are not systematically available. Definitions of unemployment also varied across surveys and are different from the modern one. That said, Eichengreen and Hatton (1988) report that more than a third of males who were looking for work in 1930 had been unemployed for at least 14 weeks and 55 percent of ongoing unemployment spells had lasted at least six months in 1940. Eichengreen and Hatton also reproduce data from Woytinsky (1942), showing the year-to-year changes in unemployment duration in Philadelphia during the 1930s. In 1933, for example, over 80 percent of the unemployed had spells of at least six months. Chatterjee and Corbae (2007) describe a special January 1931 census of the unemployed in Boston, New York, Philadelphia, Chicago, and Los Angeles, which reported that 45 percent, 61 percent, 45 percent, 61 percent, and 33 percent were jobless for at least 18 weeks, respectively.

⁵As can also be seen in figures 1 and 2 (p. 30), it took particularly long for average and median unemployment duration and the share of the long-term unemployed to return to pre-recession levels following the 1990–91 and 2001 recessions.

⁶The U-6 rate, available since 1994, includes marginally attached workers and part-time workers who want and are available for full-time work but had to settle for a part-time schedule for economic reasons. The U.S. Bureau of Labor Statistics classifies individuals as “marginally attached” if they “indicate that they want and are available for a job and have looked for work sometime in the recent past” but are not currently looking. We derived a simulated U-6 series from 1978 onward based on similar questions in the CPS. The simulated series replicates the actual reported series from 1994 onward.

⁷Okun’s law simply states a linear negative relationship exists between economic activity (that is, GDP growth) and the unemployment rate.

⁸To be clear, there are other series that are hard to forecast within this simple statistical model. We also underpredict the increase in those who are part-time workers for economic reasons and the fraction of the population outside of the labor force but not marginally attached. These results are not reported but available upon request.

⁹All of the inferences here are the same if the base of comparison is the full population rather than the labor force.

¹⁰This is true even when controlling simultaneously for all of the characteristics listed in table 1 (p. 35) in a regression framework

¹¹See Aaronson and Sullivan (1998) for similar results on job displacement and job insecurity.

¹²To implement the Blinder/Oaxaca decomposition, a separate regression is run for each time period. The change in average duration of unemployment over the two periods is then decomposed into a portion due to changes in the *levels* of the explanatory variables (for example, the fraction of females and the fraction that has completed less than high school), a portion due to changes in the *coefficients* on these explanatory variables, and a residual term that captures the effects of the *interactions* (that is, simultaneously changing the levels and coefficients).

Specifically, let unemployment duration D_{it} be specific to an individual i and a time period t . To keep things simple, we use two time periods—the 1980s, which is indexed as $t = 1$, and the 2000s, which is indexed as $t = 2$. We show the results by comparing expansions (1985–86 versus 2005–06 in the first column of table 2 on p. 36) and comparing periods of high unemployment (first half of 1983 versus second half of 2009 in table 2, second column). Duration is determined by characteristics X_{it} (for example, gender and age) that are also specific to individual i and time period t .

We can write this statistical model as $D_{it} = X_{it}b_t + e_{it}$, where e_{it} is an error term. The decomposition is then $D_1 - D_2 = (X_1 - X_2)b_2 + X_2(b_1 - b_2) + (X_1 - X_2)(b_1 - b_2)$. The first term after the equal sign is reported in the first set of rows in table 2 (“due to changes in composition”). The second term is reported in the second set of rows (“due to changes in coefficients”), and the third term is the row labeled “interactions between changes in composition and coefficients.”

Running this decomposition on the share of the unemployed undergoing long-term spells of unemployment yields similar results. Those are available upon request.

¹³Notably, changes in industrial structure have little impact. See, for example, Rissman (2009), Valletta and Cleary (2008), and Aaronson, Rissman, and Sullivan (2004) on the role of sectoral reallocation on labor market conditions during recent recessions.

¹⁴Movements between being in and out of the labor force play a much smaller role in explaining shifts in the unemployment rate, so this discussion largely abstracts from these transitions for simplicity. But we return to transitions between being unemployed and out of the labor force (*UO* and *OU*) during the most recent recession later in the article.

¹⁵The term “separations,” however, is often used elsewhere to represent all transitions out of a particular job, including job-to-job transitions. The separation and hiring rates reported in the U.S. Bureau of Labor Statistics’ *Job Openings and Labor Turnover Survey* (JOLTS) and *Business Employment Dynamics* (BED) survey also include out of the labor force transitions.

¹⁶Mazumder (2008), using the U.S. Census Bureau’s *Survey of Income and Program Participation* (SIPP), also finds that the separation rate has been of somewhat greater importance in recent recessions than suggested by Shimer (2007).

¹⁷Typically, *EE* is a continuously employed person. However, it can also be someone who transitions from one job to another without a spell of non-employment.

¹⁸It is important to note that there is an “adding up” constraint because the three probabilities must sum to 1. Therefore, it is not possible to vary the paths of all three variables simultaneously.

¹⁹It should be noted that this transition model is not additive. Allowing all of the transitions starting from *E* and all the transitions starting from *U* to follow their actual course (simultaneously) accounts for about 3.4 percentage points of the actual increase of 5 percentage points in the unemployment rate, leaving some significant share of the increase attributable to changes in transitions starting from *O*.

²⁰We also match the rise in the share of the long-term unemployed quite well.

²¹While this result is unlikely to be surprising, it should be noted that it need not be the case. The length of unemployment can increase, with a lag, from a surge in job separations.

²²Some of the variation in federal extensions occurs at the state level because of state-specific triggers for unemployment insurance benefit extensions that depend on the severity of unemployment at the state level.

²³Specifically, for each month we multiply the difference in the maximum eligibility of UI benefits over and beyond 26 weeks by the elasticity of a one-week increase in extensions on average

duration of unemployment. This product is scaled by the fraction of the unemployed receiving UI benefits in that month. The resulting estimate represents the full effect of the extension over some period of time. We then divide this effect by 12 to effectively spread out the total effect over the next 12 months. Finally, we take a running sum of the effects over the previous 12 months. Starting in December 2008, when maximum UI eligibility exceeded one year, we began to spread the effect over the next two years. See note 24 for more details on the choice of how long to spread out the impact of the extension.

²⁴The effect of an extension on the average duration of unemployment is not instantaneous. For example, the elasticity of 0.2 from Katz and Meyer (1990) is based on simulating their model on individuals over a two-year period. They found similar results from simulating the model over one year or three years. If we were to spread the effect of the initial increase in benefits over two years, this would lower the estimated contributions of the UI extensions only up until October 2009, but would have no effect on the total contributions as of December 2009.

²⁵Note that there is no spike at 26 weeks, the typical maximum number of weeks of UI eligibility. Although the CPS does not show a spike, administrative unemployment insurance records typically do (see, for example, Ham and Rea, 1987; Katz and Meyer, 1990; and Meyer, 1990, 1995).

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Do labor market activities help predict inflation?

Luojia Hu and Maude Toussaint-Comeau

Introduction and summary

Price stability is an important element in maintaining a healthy economy. Volatile prices, especially when unanticipated, can have a negative impact on aggregate demand, as people are not able to adjust and protect the real value of their financial wealth.¹ Such uncertainty can result in disruptions in business planning and reductions in capital investment spending, which could be detrimental for the long-run growth potential of the economy. In addition, inflation can impact economic welfare as wealth and income redistributions occur among different agents (Doepke and Schneider, 2006; and Franke, Flaschel, and Proano, 2006).

As experiences of some Latin American countries with hyperinflation have shown, economic growth can be seriously impaired by very high inflation (Heyman and Leijonhufvud, 1995; and Rogers and Wang, 1993). But even at much less severe levels, inflation matters. The U.S. recessions of 1973–1975, 1980, and 1981–82 were all preceded by elevated levels of inflation (Gordon, 1993).

Because of the intrinsic role of price stability in a healthy economy, controlling inflation is a primary objective of monetary policymakers. Understanding the nature of business cycles and short-run inflation dynamics is essential for the appropriate conduct of monetary policy (Svensson, 1997; and Clarida, Galí, and Gertler, 2000). In order to control inflation effectively, policymakers need to identify key indicators that help to predict inflation. Among these factors, labor market activities and, in particular, wages are closely watched. Indeed, since Phillips' (1958) paper demonstrated that there is an inverse relationship between the rate of change in money wages and the rate of unemployment, the relevance of the labor market and, in particular, the link between wages and prices have been taken as given, as noted in Fosu and Huq (1988).

It is unclear whether wage inflation causes price inflation or vice versa. If rising demand for goods and services reduces unemployment (causing it to fall below some natural rate), inflationary pressures might develop as firms bid against each other for labor and as workers feel more confident in pressing for higher wages. Then higher wages could lead to still higher prices. (In an extreme case, this might lead to a wage–price spiral, which we saw in the United States during the 1970s [Perry, 1978]).

However, if rising demand for goods and services (for example, too much money chasing too few goods) induces firms to raise their prices, these price increases and greater profits could entice workers to demand higher wages. In such an environment, inflation could lead to wage growth (Friedman, 1956; Cagan, 1972; and Barth and Bennett, 1975).

If productivity growth drives higher wages, the firm does not have to pass on higher wages into higher prices. Increased productivity therefore should curb inflationary pressures.²

A large body of research has aimed to model the inflation process empirically. However, as a recent review indicates, there is no consensus view of the best explanation for inflation (Rudd and Whelan, 2005). The literature focusing on how the labor and product markets interact has also produced mixed results. Much of the evidence suggests that wage growth, even adjusted for productivity, is not a causal factor in determining price inflation. However, inflation does

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help predict wages (Mehra, 1991, 1993, and 2000; Huh and Trehan, 1995; Emery and Chang, 1996; Hess, 1999; and Campbell and Rissman, 1994).

In this article, we revisit this question by conducting an empirical analysis of the role of labor market activities in inflation, including an examination of the relationship between productivity-adjusted labor costs (unit labor costs), unemployment, and price inflation. We contribute to the body of existing evidence with our use of updated and more recent data, including data for the past ten years. After incorporating alternative empirical approaches and elements from previous studies, we reach a fairly simple conclusion. Wage inflation is not very informative for predicting price inflation, especially during the period from 1984 onward, which has been dubbed by economists as “the Great Moderation.” However, price inflation does seem to help predict wages. We find that the unemployment data contain additional information for both wages and prices, which supports a Phillips curve type of relationship between them (Stiglitz, 1997).

In the next section, we provide a brief review of the theoretical and empirical approaches to modeling price and wage inflation. Then we present our data and discuss the econometric model of the wage and price relationship. Finally, we test for the direction of causality between wages and prices.

Theoretical background: Modeling inflation

Irrespective of the causes for inflation, the tight relationship between wages and prices follows the paradigm of the profit-maximizing firm. In its simplest form, the firm hires labor until the cost of hiring one additional worker equals the revenue that she generates. The cost of an extra labor unit (worker) is taken as the going wage rate (assuming that workers are homogeneous and the firm hires in a spot labor market, where transactions happen immediately). The firm sells its product in a spot market. The additional revenue that the firm gets from hiring one additional worker is equal to the market price of the product times the extra output that she produces. In such a market, the output price is determined by the price of the labor inputs and their productivity. This implies productivity-adjusted nominal wages grow at the same rate as product prices. In this simplified world, where the firm is a price taker in labor and product markets, the price inflation–wage inflation gap is always equal to zero.

If these assumptions are relaxed, some conditions that arise can weaken the tight link between wages and prices in the short run. As discussed in Campbell and Rissman (1994) and Huh and Trehan (1995),

labor market imperfections and certain frictions, such as adjustment costs (for example, the cost of changing employment or the presence of nominal wage rigidities), can create a wedge between the marginal product of labor and the wage rate. Such a wedge would weaken the simple framework’s strong connection between price and wage inflation. In this case, in the short run there would be a deviation away from the long-run equilibrium between price inflation and productivity-adjusted wage growth. However, over time, price and wage inflation should revert to their equilibrium relationship.

The original Phillips curve model was formulated as a wage equation relating wage inflation to the unemployment gap. But the idea that systematic movements in prices and wages may be correlated is linked to the rationalization of other formulations of the model, such as the expectation-augmented Phillips curve. Attributed to versions of work by Robert J. Gordon and also known as the Gordon triangle model of inflation, the expectation-augmented Phillips curve suggests that prices are set as a markup over productivity-adjusted wages and are affected by cyclical demand dynamics, such as unemployment gaps or output gaps, and supply shocks, such as oil price shocks. In turn, wages are a function of expected prices and demand and supply shocks. Expected prices depend on past prices (Gordon, 1982, 1985; and Stockton and Glassman, 1987). The Gordon triangle model implies a relationship between wages and prices that runs in both directions in the long run. If the proposition is correct (and assuming the markup is constant or slow-moving), then long-run movements in prices and labor costs are correlated. In the short run, if prices are slow to respond to shocks in the labor market (and we allow for short-term dynamics in such behavior), we should also further expect that short-run movements in labor costs would help predict short-run movements in prices. A number of previous researchers have sought to establish the direction of causation between wages and prices, using the framework of the expectation-augmented Phillips curve and Granger causality tests³ (Mehra, 1991, 1993, 2000; Huh and Trehan, 1995; Gordon, 1988, 1998; Emery and Chang, 1996; Hess, 1999; Campbell and Rissman, 1994; and Ghali, 1999).⁴

As noted by Stock and Watson (2008, p. 1), the traditional backward-looking Phillips curve “continues to be the best way to understand policy discussions about the rates of unemployment and inflation.” Much of the evidence in the empirical literature based on the backward-looking Phillips curve suggests that wages are not a causal factor in determining inflation. However, price inflation does help predict wages.

As in much of the literature, Campbell and Rissman (1994) and Mehra (2000) find that wages do not help predict prices. Ghali (1999), a rare exception, finds that they do. In terms of econometric methodology, all three papers include an error correction (EC) term in the Gordon triangle model to accommodate some nonstationarity in the series and allow for co-integration (the long-run relation between prices and wages). Once they establish a relationship, they test for the direction of the causality via Granger causality tests. The measures of prices and wages in the three papers are similar. Prices are measured by the gross domestic product (GDP) price deflator, and productivity-adjusted wages are measured by unit labor costs.

In these papers, the authors consider different sample periods. Campbell and Rissman (1994) use data that cover 1950:Q1–1993:Q3; Mehra (2000) uses data for 1952:Q1–1999:Q2 and considers some subsample periods; and Ghali (1999) uses data covering 1959:Q1–1989:Q3. The analyses also differ in the way they transform the data to ensure stationarity.⁵ As we mentioned previously, to accommodate the co-integration relation of the time series, all three papers use error correction models (ECM); however, Campbell and Rissman assume a known (one-to-one) error correction, while Mehra and Ghali assume that the EC is unknown and estimate the co-integration equation. The papers differ in the ways they capture short-run dynamics of supply and demand factors. Campbell and Rissman consider only demand, which they proxy with an unemployment rate variable. Mehra also uses only demand, but proxies it with the output gap and changes in unemployment rates. Ghali uses both demand (output gap) and supply (relative import prices). The papers also differ in their assumptions regarding the exogeneity of the demand and supply variables. Campbell and Rissman and Mehra assume that the demand factors are exogenous in the long-run equilibrium relation, so their variables do not enter the co-integration equation. But Ghali allows both demand and supply variables to enter the co-integration equation. Finally, they use different estimation methods. Campbell and Rissman use ordinary least squares (OLS). Mehra also uses OLS, but includes a first-step estimation of the co-integration equation between prices and wages. Ghali uses full maximum likelihood estimation (MLE), a technique that allows for multiple co-integration equations among prices, wages, and the demand and supply variables.

In this article, we incorporate various elements of these three papers to conduct (in-sample) forecasting of wage and price inflation within an expectation-augmented Phillips curve framework. To be consistent

with the literature that suggests that the time period matters, we conduct the analysis on both a full sample (which includes updated data for the past ten years), 1960:Q1–2009:Q2, and a subsample, 1984:Q1–2009:Q2. We then conduct in-sample causality tests of several versions of the error correction model: 1) assuming a known versus an unknown co-integration relation; 2) including both supply shocks and demand dynamics with alternative measures; and 3) treating supply shocks and demand as exogenous versus endogenous.

A number of studies have looked into a new version of the Phillips curve model, the so-called new Keynesian Phillips curve, or NKPC (Chadha, Masson, and Meredith, 1992; and Fuhrer and Moore, 1995); however, this approach is not within the scope of our work in this article. This new model emphasizes staggered (spread out over time) nominal wages and assumes price setting by forward-looking agents. The main difference between the traditional Phillips curve and the NKPC is that in the latter, expected future inflation is the determinant of current inflation, whereas in the traditional expectation-augmented Phillips curve, lagged inflation plays a major role. As formalized in Yun (1996), the Calvo (1983) model of staggered pricing and the Taylor (1980) model of staggered contracts are the workhorses of the NKPC. For example, Galí and Gertler (1999) and Mehra (2004) use a specification of the NKPC inflation model in which current inflation is modeled as a function of contemporaneous demand factors and of both lagged and expected inflation. Sbordone's (2002) model also emphasizes staggered nominal wage and price setting by forward-looking agents, but allows for imperfect competition with nominal price rigidity, implying an equilibrium pricing condition whereby current inflation is linked to lagged inflation and expected future real marginal costs. In sum, the main differences among these different studies are both the degree to which forward-looking, as opposed to backward-looking, elements matter and the way in which the inertia in prices is introduced (Calvo prices versus Taylor contracts).

Data

As a starting point, we take a look at the data on wages and prices and the other demand and supply economic indicators for our sample period, 1960:Q1–2009:Q2. We define prices as the GDP deflator consistent with the three papers we discussed earlier—Campbell and Rissman (1994), Mehra (2000), and Ghali (1999).⁶ For wages, we use unit labor costs for the nonfarm business sector (ULC). ULC is nominal wages, adjusted for labor productivity ($ULC = W \times L/Y$, where W equals nominal wages, L equals hours per

BOX 1**Definitions of variables**

$p = \log(\text{GDP deflator})$, where GDP is gross domestic product

$w = \log(\text{ULC})$, where ULC is unit labor costs for the nonfarm business sector

$\pi^p = \Delta p$, quarter-to-quarter growth rate of GDP deflator

$\pi^w = \Delta w$, quarter-to-quarter growth rate of ULC

$g = \log(\text{real GDP/potential GDP})$; that is, the output gap

$u = \text{unemployment rate} - \text{nonaccelerating inflation rate of unemployment (NAIRU)}$; that is, the unemployment gap

$imp = \log(\text{relative import price deflator inclusive of oil/GDP deflator})$

worker, and Y equals output, implying $ULC = W/(Y/L)$. Box 1 summarizes the definition of the variables used in this analysis.

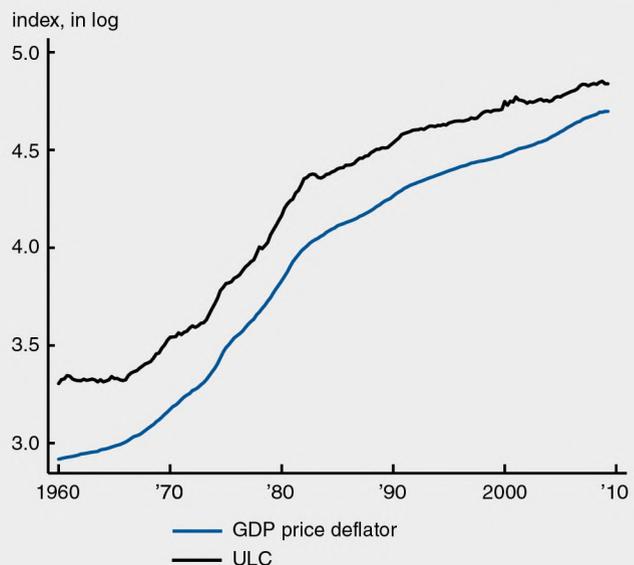
Figure 1 charts the time series of the GDP price deflator and ULC over the period 1960:Q1–2009:Q2. This chart clearly shows the correlation between the two series.

In figure 2, we report the quarter-to-quarter change (annualized) in the two series.⁷ We note two distinctive periods: Inflation and wage growth increased in quite dramatic fashion in the 1970s (this is the period known for the wage–price spiral phenomenon). From the mid-1980s onward, we see a tapering off of inflation and wage growth.⁸ Looking more closely at the co-behavior of the two series, from the mid-1960s up to 1984, the two series show quite a lot of co-movement. From 1984 onward, there appears to be much less co-movement between wage growth and price inflation. In fact, while wage growth continues to fluctuate, price inflation remains markedly low and stable. This figure suggests that the relationship between the two series may not be stable over the full sample period and that, as others analyzing trends in inflation and wage growth have suggested, these series may not have a “normal,” or built-in, level and therefore shocks to them could be quite persistent (Fuhrer and Moore, 1995; and Benati, 2008).

The difference between the quarter-to-quarter inflation rate of the GDP deflator (π^p) and growth rate of ULC (π^w) is shown in figure 3. The difference can be viewed as representing a deviation from the long-run equilibrium (assuming a one-to-one or unit relationship, $EC = \pi^p - \pi^w$). This is clearly a simplifying assumption. We later consider versions of the model that assume constrained co-integration, where we impose unit coefficients, but we also consider a version of the model with unconstrained co-integration, where we estimate the coefficients for the error correction term.

Following the theoretical proposition of the profit-maximizing firm, such a deviation should revert to its mean in the long run. Consistent with this, in figure 3 we note that the disequilibrium term has been fluctuating around a mean of zero (that is, it has not gone up or down over time in a discernible trend). There is clearly a long-term relation between the two series, but it is unclear whether there is a causal relationship or, if there is, which one causes the other.

Figures 4 and 5 report the measures of excess demand or slack in the economy—that is, the unemployment gap and output gap. As noted in box 1, the unemployment gap is the difference between the civilian unemployment rate and the nonaccelerating inflation rate of unemployment (NAIRU). The NAIRU is provided by the Congressional Budget Office (CBO),

FIGURE 1**Prices and unit labor costs, 1960:Q1–2009:Q2**

Note: For further details on the gross domestic product (GDP) price deflator and unit labor costs (ULC), see box 1.

Source: Authors' calculations based on data from the U.S. Bureau of Economic Analysis from Haver Analytics.

and it is an equilibrium rate that does not tend to increase or decrease the inflation rate. The output gap is the logarithm of the ratio of real GDP to potential real GDP. Potential real GDP is also estimated by the CBO. As can be expected, we note in these figures that unemployment increased and the output gap decreased in periods of economic slowdown (for example, in the 1970s, 1980s, and early and late 2000s).

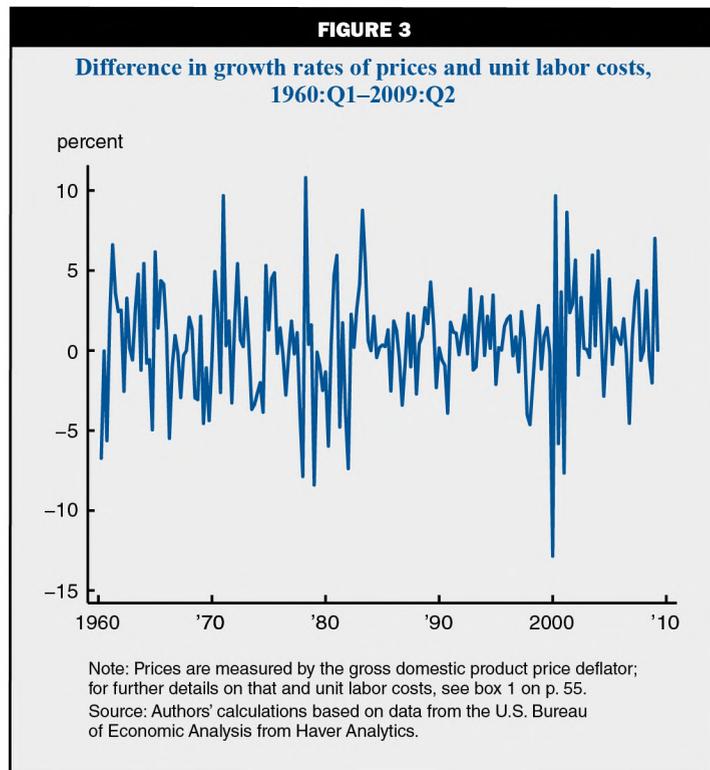
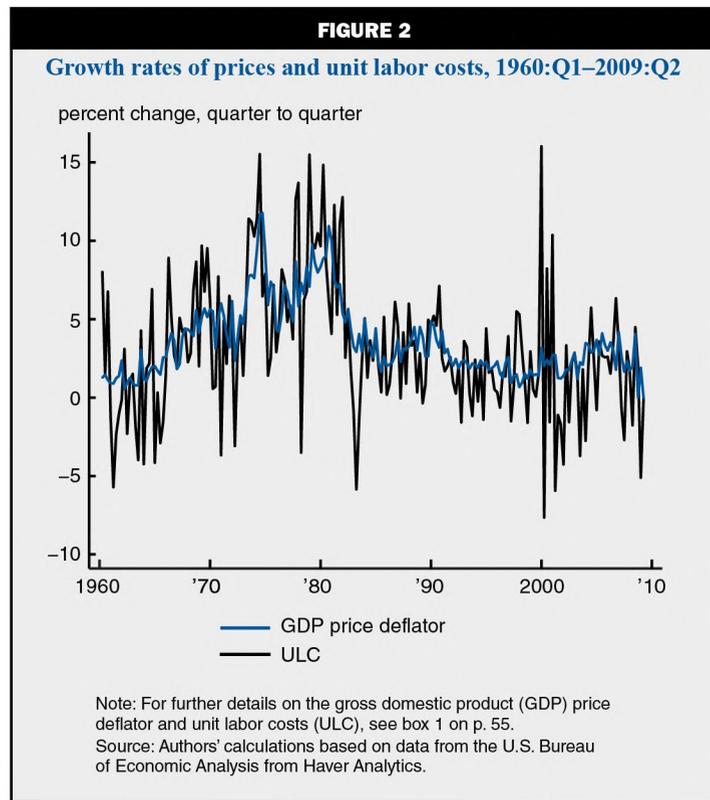
Finally, figure 6 shows the time series of the relative prices of imports, a measure of supply shocks. The role of import prices is fairly obvious. The aggregate supply curve should shift when input prices change, and input prices are affected by the prices of imports. The figure shows that prices of imports changed very little in the 1960s and early 1970s. They increased substantially in 1974 and again in 1979–80. Since 1981, relative import prices have changed very little. We would therefore expect that this variable should be relatively less important for explaining inflation in the past three decades.

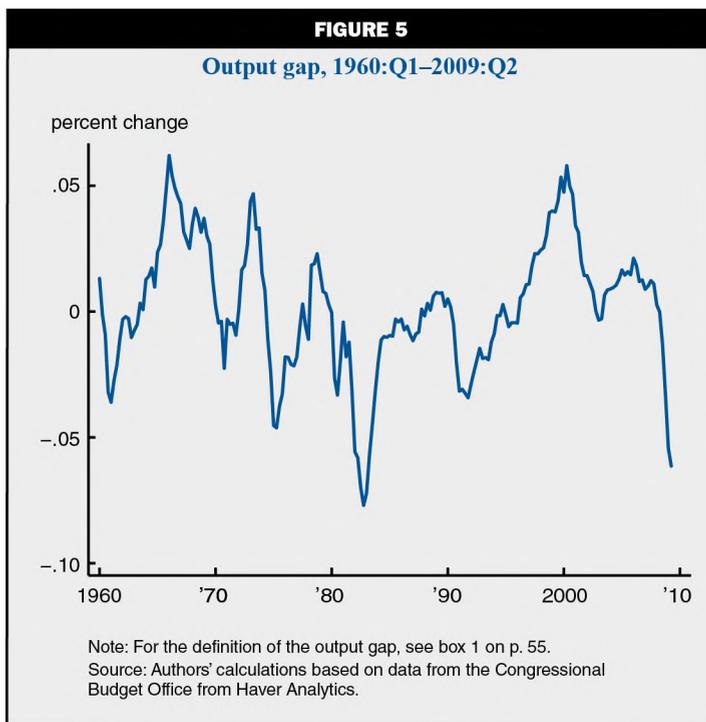
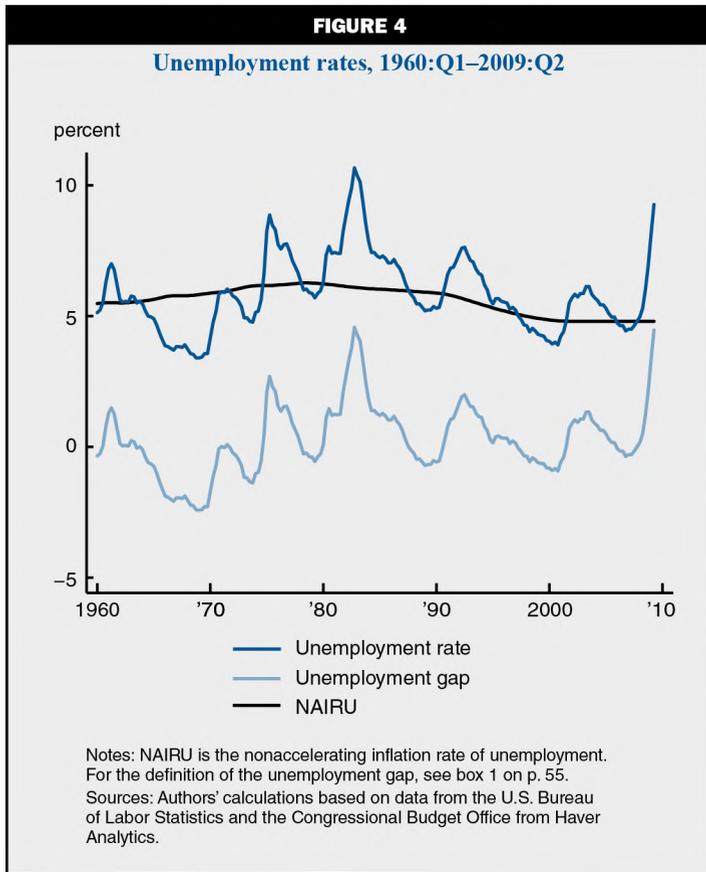
Empirical estimation

To make clear the hypotheses that we will be testing, it is useful to describe in more specific terms the expectation-augmented Phillips curve model. The basic relationships are represented by the following system of equations:

- 1) $\pi_t^p = h_0 + h_1 \pi_t^w + h_2 DD_t + h_3 SS_{pt}$,
- 2) $\pi_t^w = k_0 + k_1 \pi_t^{e,p} + k_2 DD_t + k_3 SS_{wt}$,
- 3) $\pi_t^{e,p} = \sum_j \lambda_j \pi_{t-j}^p$.

where π_t^p is the first difference of the log of the price level; π_t^w is the first difference of the log of the nominal rate of ULC; DD_t is a vector of demand pressure variables, which include g (the output gap) and/or u (the unemployment gap) as defined previously.⁹ The term $\pi_t^{e,p}$ is the expected inflation level, SS_{pt} represents supply shocks affecting the price equation, and





SS_{wt} represents supply shocks affecting the wage equation. Such supply shocks are proxied by imp (the relative import prices inclusive of oil) and two period dummies indicating President Nixon's price and wage control periods. (The first period is 1971:Q3–1972:Q4, and the second period is 1973:Q1–1974:Q4.)

As can be seen, equation 1 reflects the idea that prices are a markup over productivity-adjusted wages and are affected by cyclical demand and relative supply shocks. Equation 2 shows that wages are affected by demand and supply and expected price level. Equation 3 shows that expected inflation is a function of past prices. Further, to accommodate the statistical features of the time series, we include an error correction term in the Gordon triangle model (equations 1–3). We also keep the demand and supply variables to affect the short-run dynamics of prices and wages. This is represented as follows:

$$\begin{aligned} \Delta \pi_t^p &= \alpha^1 (\pi_{t-1}^p - \pi_{t-1}^w) + \sum_{l=1}^L \gamma_l^1 \Delta \pi_{t-l}^p \\ &+ \sum_{l=1}^L \lambda_l^1 \Delta \pi_{t-l}^w + \sum_{l=1}^L \rho_l^1 DD_{t-l}^p \\ &+ \sum_{l=1}^L \phi_l^1 SS_{t-l}^p + \varepsilon_t^1, \end{aligned}$$

$$\begin{aligned} \Delta \pi_t^w &= \alpha^2 (\pi_{t-1}^p - \pi_{t-1}^w) + \sum_{l=1}^L \gamma_l^2 \Delta \pi_{t-l}^p \\ &+ \sum_{l=1}^L \lambda_l^2 \Delta \pi_{t-l}^w + \sum_{l=1}^L \rho_l^2 DD_{t-l}^w \\ &+ \sum_{l=1}^L \phi_l^2 SS_{t-l}^w + \varepsilon_t^2. \end{aligned}$$

The error correction term ($EC = \pi_{t-1}^p - \pi_{t-1}^w$) allows for a long-run equilibrium relationship between price and wage inflation. The parameter α therefore reflects long-run dynamics, and γ and λ capture short-run dynamics. The term ε_t^1 is the residual from the price equation, while ε_t^2 is the residual from the wage regression. L is the maximum number of lags on the various variables needed to make the random disturbances serially uncorrelated. Again, as previously noted, DD and SS are vectors of variables representing

demand and supply shocks affecting price and wage inflation, as in some previous studies (for example, Mehra, 2004; and Hess and Schweitzer, 2000).

We have the following hypotheses concerning the joint short-run and long-run equilibrium relationships in wages and prices: Hypothesis 1 is that wages do not predict prices ($H_0 : \alpha^1 = 0, \lambda_1^1 = 0, \dots, \lambda_L^1 = 0$). Hypothesis 2 is that prices do not predict wages ($H_0 : \alpha^2 = 0, \gamma_1^2 = 0, \dots, \gamma_L^2 = 0$).

Recalling that the parameter α reflects long-run dynamics, while γ and λ capture short-run dynamics, we test for the hypotheses and determine the sources of the short-run and long-run co-movements between wages and prices, using Granger causality tests. Our test for Granger causality involves examining whether lagged values of one series (that is, wages) have significant explanatory power for another variable (that is, prices). In this exercise, both variables may Granger-cause one another.¹⁰ Both series in question may also be co-integrated.

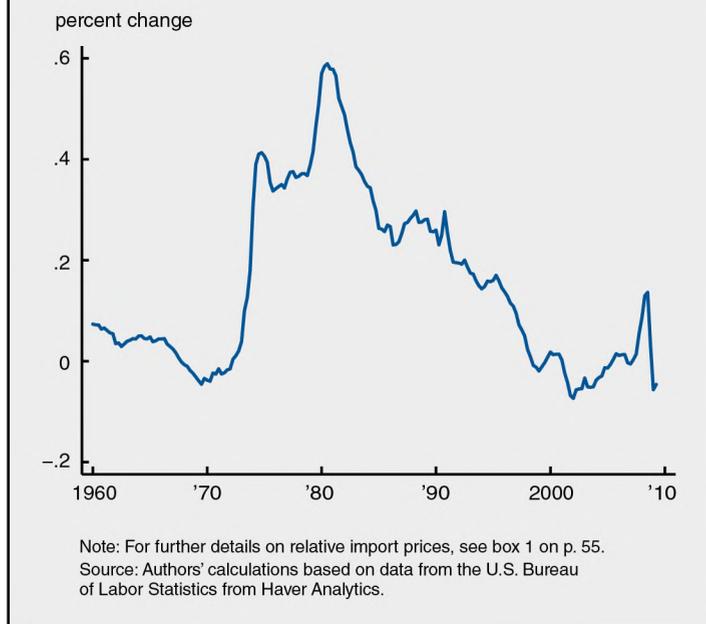
Recall that by incorporating an error correction term in the Granger causality tests, we allow the series in levels to catch up with or equal one another. The significance of the error correction term in the Granger causality test would signal the fact that the series in question are driven to return to a long-run equilibrium relationship that is causal.

Granger causality test results

Before conducting the Granger causality tests, we examined the stationarity of the series. The results of the augmented Dickey–Fuller (ADF) unit root tests for price inflation and wage inflation confirmed that we cannot reject the null hypothesis of a unit root at the 1 percent level—that is, the growth rates of prices and wages are both integrated of order one, $I(1)$. Also, for the full sample period (1960:Q1–2009:Q2), the relative import prices (*imp*), unemployment gap (*u*), and output gap (*g*) are also all $I(1)$.

Table 1 presents the results of our tests for Granger causality between wages and prices for the full sample period, 1960:Q1–2009:Q2, and for a subsample period, 1984:Q1–2009:Q2. In this bivariate model, we assume that $DD = 0$ and $SS = 0$. The regression includes lagged prices and lagged unit labor cost growth. The number of lags for each variable is set to four ($L = 4$). Panel A of table 1 reports the evidence on whether the column variables Granger-cause price inflation, while panel B

FIGURE 6
Relative import prices, 1960:Q1–2009:Q2



shows the evidence for whether the column variables Granger-cause wage growth. The error correction column refers to the long-run effect; the wages column (in panel A) and prices column (in panel B) refer to the short-run effect; and the joint hypothesis column refers to the long- and short-run effects. Each column reports the p value, the level of statistical significance with which one can reject the null hypothesis. A high p value should be taken as evidence that the column variable does *not* Granger-cause price or wage inflation.

Referring back to the ECM, to be co-integrated, at least one of the α_s in the two equations should not be equal to zero. Looking at panel A of table 1 for the full sample period, 1960:Q1–2009:Q2, the high p value in the error correction column means that $\alpha^1 = 0$. Therefore, we can say that prices do not catch up with wages in the long run. But rather wages adjust to catch up with prices ($\alpha^2 \neq 0$), per the low p value for error correction in panel B. The high p value for hypothesis 1 of the joint test of error correction and wages (panel A, third column) suggests that wages don't help predict prices in either the short run or the long run (at a 5 percent significance level).

To summarize the results in table 1, wages do not cause price inflation in our Granger causality tests. However, prices do cause wage inflation. Wages, but not prices, adjust to maintain the long-run equilibrium relationship. This is true for both the full sample (1960:Q1–2009:Q2) and the subsample (1984:Q1–2009:Q2).

TABLE 1						
Granger causality test: Bivariate model						
Period	A. Are prices caused by			B. Are wages caused by		
	Error correction	Wages	Hypothesis 1: Joint test of error correction and wages	Error correction	Prices	Hypothesis 2: Joint test of error correction and prices
1960:Q1–2009:Q2	0.32	0.34	0.06	0.00	0.06	0.00
1984:Q1–2009:Q2	0.82	0.29	0.27	0.00	0.38	0.00

Notes: The number of lags for each variable is set to four. Each column reports the *p* values, indicating the level of statistical significance for the test that the column variable does not Granger-cause either price inflation or wage inflation. See the text for details on hypothesis 1 and hypothesis 2.
Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics from Haver Analytics.

Besides the price–wage inflation gap, our model stipulated that there are other short-run demand and supply determinants of price and wage inflation. To allow for these cyclical (that is, excess) demand factors to additionally affect wages and prices in the short run, we add the unemployment gap (and, alternatively, the output gap, which we do not report in the table) to our regressions. We also add supply variables, as proxied by the relative import prices and dummy variables for the Nixon price and wage control periods. We run the regressions with these demand and supply control variables in differences, as we found that they were I(1). Both demand and supply control variables include their lags, which were set to four. And again, we include the error correction term.

Table 2 reports the *p* values from the Granger causality tests for this augmented model. The results in both panels A and B of this table suggest that wages do not predict prices; however, prices do predict wages. In the long run, wages adjust to the error correction, while prices do not. In other words, price and wage inflation move together in the long run because wages adjust to close the gap, and not because price inflation responds to wage growth.

As for the additional regressors, for the full sample, the unemployment gap has additional predictive power for both price and wage inflation, while the relative import prices only help predict price inflation. For the subsample, 1984:Q1–2009:Q2, the unemployment gap only helps predict price inflation, while the relative import prices do not help predict either. Using alternative measures of excess demand (for example, changes in the unemployment rate or the output gap) yields qualitatively similar results, which we do not report here.

We find the result of the informational content of the unemployment gap for both wage and price inflation interesting; it suggests that such cyclical variables

play an important short-term role in determining inflation (Campbell and Rissman, 1994). In the tradition of a Phillips curve type of relationship, price inflation thus appears to be still very much a labor market phenomenon (Stiglitz, 1997).

The two models that we have discussed thus far constrain the co-integration relationship between price inflation and wage inflation to be one to one, which can be justified by theory under the assumption of perfect competition and a Cobb–Douglas production function (for example, as in Campbell and Rissman, 1994). However, this might be too restrictive an assumption.

We relax this restriction and consider a generalized model, allowing for an unconstrained co-integration relationship between prices and wages (that is, in the unconstrained case, we estimate the coefficients for the error correction terms). Moreover, we also allow the supply and demand variables to enter the long-run equilibrium relation. In other words, the supply and demand variables are now treated as endogenous and could enter the error correction. (For simplicity, we do not reproduce the new augmented generalized ECM, but note that this means our model now gets augmented by two more equations with the demand and supply variables on the left-hand side). First, looking at the *p* value results for the joint short-run and long-run hypothesis between wages and prices based on the unconstrained model in table 3, panels A and B, we note that similar to the results in table 2, wages do not help predict prices, but prices do help predict wages.

Recall that in this new unconstrained model, the coefficients for all the variables are being estimated. This ECM was estimated by the maximum likelihood estimation technique. For the full sample, the model was found to be co-integrated with rank 2 (that is, it has two unique co-integration relationships). The two estimated co-integration relationships, with the standard errors in parentheses, are as follows:

TABLE 2

Granger causality test: Multivariate model

Period	A. Are prices caused by					B. Are wages caused by				
	Error correction	Wages	Unemployment gap	Relative import prices	Hypothesis 1: Joint test of error correction and wages	Error correction	Prices	Unemployment gap	Relative import prices	Hypothesis 2: Joint test of error correction and prices
1960:Q1–2009:Q2	0.26	0.63	0.00	0.00	0.29	0.00	0.00	0.00	0.09	0.00
1984:Q1–2009:Q2	0.87	0.16	0.00	0.10	0.17	0.00	0.21	0.05	0.75	0.00

Notes: The number of lags for each variable is set to four. Each column reports the p values, indicating the level of statistical significance for the test that the column variable does not Granger-cause either price inflation or wage inflation. See the text for details on hypothesis 1 and hypothesis 2.

Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics and Congressional Budget Office from Haver Analytics.

TABLE 3

Granger causality test: Multivariate model with unknown co-integration parameters

Period	A. Are prices caused by					B. Are wages caused by				
	Error correction	Wages	Unemployment gap	Relative import prices	Hypothesis 1: Joint test of error correction and wages	Error correction	Prices	Unemployment gap	Relative import prices	Hypothesis 2: Joint test of error correction and prices
1960:Q1–2009:Q2	0.25	0.40	0.00	0.00	0.06	0.00	0.00	0.09	0.01	0.00
1984:Q1–2009:Q2	0.62	0.12	0.00	0.04	0.15	0.00	0.22	0.53	0.21	0.00

Notes: The model was estimated by using the maximum likelihood estimation technique. The number of lags for each variable, chosen by the Akaike Information Criterion (AIC), is set to four. Each column reports the p values, indicating the level of statistical significance for the test that the column variable does not Granger-cause either price inflation or wage inflation. See the text for details on hypothesis 1 and hypothesis 2.

Sources: Authors' calculations based on data from the U.S. Bureau of Labor Statistics and Congressional Budget Office from Haver Analytics.

$$\pi^p = 2.84 - 1.40u + 8.27 imp,$$

(0.35) (2.35)

$$\pi^w = 1.89 - 1.98u + 11.76 imp.$$

(0.42) (2.75)

As can be seen, the unemployment gap and relative import prices variables enter both co-integration equations significantly. However, we find that the adjustment parameters on the error correction terms in the equations of the unemployment gap and relative import prices were statistically insignificant. (For simplicity, we do not report the unemployment gap and relative import prices equations here.) This suggests that these two variables do not adjust (as wages and prices do) to maintain the long-run equilibrium relations. In fact, the likelihood ratio test for the null hypothesis that the adjustment parameters in the unemployment gap and relative import prices equations are jointly zero has a p value of 0.68.

For the subsample, 1984:Q1–2009:Q2, the unemployment gap is $I(2)$ instead of $I(1)$. After replacing the unemployment gap by its first difference, the model is estimated to have one co-integration relation. In this case, the unemployment gap and the relative import prices do not even enter the co-integration equation significantly. The unemployment gap and the relative import prices appear to be exogenous in the long-run equilibrium, especially in the subsample period.

Conclusion

Much research has been devoted to not only identifying the causes of inflation but also gauging which economic indicators could best measure and predict inflation. Using more recent and updated data, we analyzed labor market indicators, namely, productivity-adjusted wages and unemployment (as well as supply shock and demand factors), to determine the extent to which they contain information to help predict inflation.

Similar to previous research, we have found that wage growth does not cause price inflation in the Granger causality sense. We found this to be particularly true for the period from 1984 onward (referred to as the Great Moderation by economists). By contrast, price inflation does cause wage growth in the Granger causality sense. Moreover, unemployment has additional predictive power for inflation for the full sample (1960:Q1–2009:Q2), as well as our subsample (1984:Q1–2009:Q2). The unemployment gap is therefore a useful indicator for inflation.

As the data indicate, in recent years wage growth has been particularly slow. Given this, some analysts think that we do not have to be overly concerned about future inflation. Our findings in this article, however, do not support the claim that slow wage growth is a harbinger of low inflation.

NOTES

¹For further discussion of the effects of inflation, see, for example, Dossche (2009).

²Federal Reserve Chairman Ben S. Bernanke (2008) noted in a speech that we are unlikely to see the 1970s type of wage–price spiral in today’s economy. Crucial productivity gains that help blunt inflationary forces were among the several factors cited. Also, inflation expectations, although somewhat on the rise, are much lower than they were in the mid-1970s.

³Granger causality is a statistical methodology for demonstrating whether a variable contains information about subsequent movements in another variable.

⁴Stock and Watson (2008) provide a survey of the literature of the past 15 years, which looks at out-of-sample forecast evaluations based on Phillips curves as well as other inflation forecasting models.

⁵Mehra (2000) and Ghali (1999) treat prices and wages as integrated of order one, $I(1)$, while Campbell and Rissman (1994) treat the growth rates of prices and wages as $I(1)$.

⁶We also conducted the analysis using the U.S. Bureau of Economic Analysis’s Personal Consumption Expenditures Price Index as the price measure. Generally, the results were similar.

⁷The results in the subsequent analysis are largely robust to an alternative measure of inflation using a four-quarter change in price.

⁸As mentioned earlier, this period has been dubbed by economists as the Great Moderation, when macroeconomic indicators were remarkably stable (see, for example, Bernanke, 2004; Kim and Nelson, 1999; McConnell and Perez-Quiros, 2000).

⁹Several explanations have been offered in the literature to motivate unemployment in a wage and price equation. Beside the Phillips (1958) underlying model of change in wages as a function of the unemployment rate, the literature of efficiency wages provides some motivation (for example, Shapiro and Stiglitz, 1984). Huh and Trehan (1995) provide a summary of the logic of the efficiency wage approach in explaining the inclusion of unemployment in a wage and price equation. Also, see Ghali (1999) and Gordon (1988).

¹⁰More specifically, the Granger causality test is a two-step regression procedure used to examine the direction of causality between two series. For example, to determine whether there is causality running from p to w , w is first estimated as a function of past values of w (this is called the restricted equation). Then w is estimated as a function of past values of w and past values of p (this is called the unrestricted regression). There is causality in the Granger sense from p to w if the inclusion of the past values of p significantly improves the estimation of w (that is, by an F test).

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