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The Unemployment Impact of the 2008 Extension of Unemployment Insurance: As High as Robert Barro Suggested?

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Link to Abstract

My calculations suggest the jobless rate could be as low as 6.8%, instead of 9.5%, if jobless benefits hadn't been extended to 99 weeks.

—Robert J. Barro, *Wall Street Journal* op-ed, August 30, 2010

The U.S. unemployment rate averaged 4.6 percent in the two years prior to the economy entering the 2008-09 recession and it more than doubled to 9.5 percent by the summer of 2010 when Robert Barro (2010) estimated how much of that increase was attributable to changes in unemployment insurance (UI) policy. His analysis suggested the recession was not nearly as severe as unemployment statistics suggested because more than half of the increase in the unemployment rate since the recession began was caused by an extension of UI. Barro believes the administration should reconsider extensions of UI benefits.

The unemployment insurance program is a perfect example of government policy that confronts the classic tradeoff between economic efficiency and equality (Okun 1975). The motivation behind the creation of the UI in 1935 was clearly a sense of compassion—to provide temporary help to those suffering income losses during a widespread economic downturn. But this compassion comes at some cost. Inefficiency can result as a consequence of subsidizing the very condition against which UI seeks to protect individuals. Economists have long accepted the theory and evidence that UI benefits are likely to cause extended durations of

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unemployment (Katz and Meyer 1990; Chetty 2008), and some have recommended that the resulting inefficiency is large enough to consider reducing the extent of UI benefits (Barro) or eliminating the current UI program altogether and replacing it with individually pre-funded savings accounts (Feldstein and Altman 1998).

The current recession, with its combination of historically long durations of unemployment (Farber 2011) and unprecedented extensions of UI benefits (U.S. Department of Labor 2011), has generated renewed interest in this policy debate. Proposals to scale back UI benefits are strengthened if recent extensions in the duration of UI benefits are found to have caused large and inefficient increases in the unemployment rate. On the other hand, if these extensions caused only modest increases in unemployment during a deep recession such as 2008-2009, a policy of limiting the extension of benefits may punish those individuals disproportionately suffering from the economic downturn while generating little gain in economic efficiency. Is Barro correct? Has UI caused large, unnecessary increases in unemployment during the Great Recession? What do the most recent data tell us about this important policy question?

Bhashkar Mazumder Estimates the Impact

A number of economists have recently attempted to measure the impact of the Emergency Unemployment Compensation program—created in the summer of 2008—on the national economy's unemployment rate (U). Most of these studies have first estimated the impact of extended benefits on the average duration of unemployment and then used that estimate to calculate how much such an increase in the duration of unemployment would raise U, ceteris paribus. Bhashkar Mazumder (2011) provided an example of this methodology in a recent article published by the Federal Reserve Bank of Chicago.

Mazumder contrasted the maximum number of weeks of unemployment insurance available during the current recession (99 weeks as of February 2011) with that of the 1982 recession when maximum coverage extended to only 55 weeks. The observed unemployment rate may overstate weakness in the labor market if the unemployed, instead of taking existing job offers, have lengthened the duration of their job search as a consequence of their ability to collect UI benefits for extended periods. Since policymakers rely heavily on unemployment statistics when formulating policy, Mazumder fears that such an increase in the unemployment rate could lead to inappropriate macroeconomic policy decisions (2011, 1).

Mazumder's method of estimating how much the extension of UI has raised U is based on the steady-state unemployment rate theory that defines the economy's equilibrium unemployment rate in terms of the rate of job separation (*s*, the fraction of the employed who are periodically separated from their jobs) and the rate of job finding (*f*, the fraction of the unemployed who find employment during the same period). In steady-state equilibrium U must equal $s/(s+f)$, and if the recent extension of the maximum duration of UI has lengthened the average duration of unemployment (and thereby reduced *f*), then extended UI has increased the economy's steady-state unemployment rate. An estimate of the magnitude of the decline in the job finding rate caused by extended UI can be used to estimate the resulting increase in U, ceteris paribus.

Mazumder obtained an initial value of *f* by assuming that the economy was in a steady state during the six months prior to the July 2008 extension of UI. The unemployment rate averaged 5.1% during that period and the average duration of unemployment was 17 weeks. A seventeen-week average duration of unemployment implies a probability of finding a job in any given week of 1/17 or about 0.06, and since the average month has 4.33 weeks, he estimated the initial steady-date monthly rate of job finding to be 0.253. This combined with U = 0.051 implies a job separation rate of 0.0136. These are the base values for *f* and *s* he used to estimate how much steady-state U has been increased as a consequence of extended UI (ibid., 2).

The next step in his analysis was to obtain an estimate of how the national average duration of unemployment has been affected by the UI benefit extensions. Mazumder used the finding by Card and Levine (2000) that in a New Jersey extended UI experiment, a one-week increase in the maximum number of weeks of UI benefit eligibility raised the duration of unemployment by 0.1 week. This estimate was applied to the increase in the number of weeks the average eligible unemployed worker could receive UI as of spring 2011, allowing Mazumder to conclude that the average duration of unemployment increased from 17 to almost 20 weeks as a consequence of extended benefits, thus lowering the rate of job finding from 0.253 to 0.22 and raising the steady state unemployment rate by 0.8 percentage points (Mazumder, 2).

As a test of robustness, Mazumder provided alternative estimates by considering other measures of the impact of a week of extended UI eligibility on the duration of unemployment. If an added week of UI benefit eligibility were to increase the duration of unemployment by 0.16 weeks, as estimated by Katz and Meyer (1990), instead of 0.1, then Mazumder's estimate of the resulting increase in U would be 1.2 points rather than 0.8. He also considered alternative values for the UI take-up rate, or the proportion of the eligible unemployed who do receive benefits. During this recent recession, the take-up rate rose from a pre-recession rate of 0.4 to 0.7. If half of this increase in the take-up rate were assumed to be due to the severity of the current recession and the other half due to the extension of UI, his estimate of the resulting increase in U would change from 0.8 to 1.1 points. (Mazumder, 3) When he combined the Katz and Meyer larger estimate of the impact of an added week of UI on the duration of U with the higher take-up rate, his estimate of the resulting increase in the unemployment rate was 1.7 points (Mazumder, 4). It is worthy to note that Mazumder's estimates are much smaller than that obtained by Barro but are consistent with other recent studies Mazumder reviewed that used both similar and alternative methodologies (ibid., 1-2).

Robert Barro Estimates the Impact

A much more damning estimate of the impact of extended UI on the unemployment rate was provided by Robert Barro (2010) in an op-ed entitled, "The Folly of Subsidizing Unemployment" published in the *Wall Street Journal*. Barro suggested that the June 2010 unemployment rate of 9.5% would have been 2.7 percentage points lower had it not been for the expansion of UI benefit eligibility.

Barro's methodology is straightforward. He begins, like Mazumder, by noting a stark contrast between unemployment durations in the current recession with those of 1982 when the peak mean duration of unemployment was 21.2 weeks and 24.5% of the unemployed were experiencing joblessness for longer than 26 weeks. In contrast, the duration of unemployment in October of 2009 averaged 27.2 weeks, with 36% of the unemployed out of work for more than 26 weeks. Barro notes that since the end of WWII, the duration of unemployment has averaged less than 21 weeks and the long-term jobless made up less than 25% of the unemployed. Why are the numbers for the current recession so much higher than those in the historical record? "[T]he dramatic expansion of unemployment-insurance eligibility to 99 weeks is almost surely the culprit," says Barro.

Barro calculates the possible impact of extended unemployment duration on the unemployment rate by first assuming no change in the size of the labor force. He also assumes that the long-term share of the unemployed would have been 24.5% (its July 1983 level) instead of 46.2% (the June 2010 level) had it not been for the expansion of UI. Given these assumptions, 4.2 million persons who would have otherwise been long-term unemployed would instead have taken a job and the unemployment rate would have been 6.8% instead of 9.5%. Barro goes on to conclude that President Obama should blame his economic advisors for not "arguing that a reckless expansion of unemployment-insurance coverage to 99 weeks was unwise economically and politically."

While Barro admits that his estimate of the impact of extended UI on long-term unemployment and the unemployment rate is "rough," it is worth noting that his conclusions are based on two critical assumptions. The first involves the impact of UI on the size of the labor force. In his hypothetical absence of the expansion of UI coverage to 99 weeks "the labor force still equaled the observed value (153.7 million)." Under this assumption, U falls with the elimination of extended UI coverage because 4.2 million long-term jobless persons now take jobs they otherwise would have turned down.

Barro does not consider the possibility of an alternative mechanism whereby the elimination of extended UI benefits lowers U because jobless workers who would have otherwise continued their job search (and be counted as unemployed) while collecting extended benefits instead leave the labor force as discouraged workers. In this scenario, reducing UI benefits does lower U officially, but does not in fact increase the absolute number of employed persons.

Barro's interpretation of the impact of UI is focused exclusively on the first of these two mechanisms, and it is an interpretation that strengthens his position that extending UI benefits was inefficient because it reduces employment by over 4 million persons. If, on the other hand, extended UI raises U because it encourages the jobless to stay in the labor force, then the policy's economic impact is primarily one of an income transfer to the long-term unemployed that creates no real job losses. Recent analysis of the unemployed in Austria and a review of the existing literature covering other counties (Card et al. 2007) contradict Barro's interpretation, as, when their unemployment benefits expire, the majority of the unemployed leave the labor force rather than take a job. A recent working paper using data from the Current Population Survey (Rothstein 2011) came to similar conclusions. The impacts of extended UI benefits on unemployment exit, reemployment, labor force exit hazard rates, and on the unemployment rate were estimated. Unemployment insurance extensions were found to have had a larger impact on labor force exit than on reemployment. The December 2010 unemployment rate was estimated to be only 0.3 to 0.6 percentage points higher as a consequence (ibid., 3-4).

In addition to his assumption that extended UI benefits reduce employment rather than change job seekers into discouraged workers, Barro also assumed that the severity of the recent recession could not be responsible for extended durations of unemployment.

> The administration has argued that the more generous unemployment-insurance program could not have had much impact on the unemployment rate because the recession is so severe that jobs are unavailable for many people. This perspective is odd on its face because, even at the worst of the downturn, the U.S. labor market featured a tremendous amount of turnover in the form of large numbers of persons

> hired and separated every month. […] A program that reduced incentives for people to search for and accept jobs could surely matter a lot here. (Barro 2010)

This argument—that the severity of the Great Recession cannot explain the current long average duration of unemployment—brings us to the focus of the present investigation. In a severe recession where UI benefits have been extended by unprecedented lengths, how can we determine how much of the observed increase in the unemployment rate is due to the impact of extended UI as opposed to the downturn's real impact on job separation and potential job offers?

Recent analysis at the National Bureau of Economic Research used Displaced Workers Survey data to conclude that the 2008-09 recession did experience unusually high job separation rates compared to previous recessions, including that of 1982 (Farber 2011). While the NBER research made no attempt to estimate the impact of the increase in job separation on the unemployment rate, it strongly suggests that any attempt to measure the impact of UI changes must take into account the separate impact of the severity of the recession. The next section of the paper attempts to do just that. Two detrending methods are applied to annual real GDP and annual unemployment rate data in an attempt to quantify the separate influences of the severity of the recession and the extension of UI benefits on changes in the unemployment rate. The first detrending method—a variation of Okun's Law—explains annual changes in the unemployment rate based on deviations in the growth rate of real GDP. The second method uses the Hodrick and Prescott filter to remove the trend from annual real GDP and unemployment rates to identify their business cycle components. The resulting deviation of real GDP from the HP-filter trend is used to explain cyclical deviations in the unemployment rate. Variables measuring the duration of extended UI benefits are then included in both models to measure the impact of extended UI on the unemployment rate after controlling for the severity of the recession.

Theory and Evidence

Okun's Law is a rule of thumb describing the historically strong statistical relationship between changes in the unemployment rate and the growth rate of real GDP. Arthur Okun (1962) used quarterly data for the 1947-1960 period to estimate the "first differences"[2] version of the relationship

2. Okun also defined a "gap" version of the relationship between unemployment and real output by expressing changes in the unemployment rate as a function of the gap between potential and actual output.

change in the unemployment rate = a + b*growth rate of real GNP

The focus of Okun's work was to identify a time series for potential GNP (1962, 2), but what came to be known as Okun's Law was his statistical documentation of the unemployment-output growth rate relationship (Knotek 2007). That relationship is used in this section to compare observed changes in the unemployment rate during the recent recession with increases that would be predicted based on the historical relationship between GDP growth and unemployment. Regression analysis is used to estimate the impact that the extension of UI has had on the unemployment rate after controlling for the state of the macroeconomy as measured by the annual growth rate of real GDP.

I used annual observations on the growth rate of real GDP (*grgdp*) (U.S. Bureau of Economic Analysis 2011) and the annual change in the unemployment rate for all civilian workers (Δu) (U.S. Bureau of Labor Statistics 2010) for the 1960-2010 period to obtain an OLS estimate of the "first differences" version of Okun's Law, $\Delta u = \alpha + \beta * grgdp$. The regression results are shown in equation 1 with the T ratios in parentheses. The R-squared for the estimated equation was .749.

$$\text{Equation 1: } \Delta u = \underset{(10.4)}{1.313} - \underset{(-11.98)}{0.392} * grgdp$$

The intercept term in Equation 1 provides an estimate of how much the unemployment rate would change in a year of no GDP growth, and the slope coefficient estimates the impact of a one percentage point increase in the GDP growth rate on the annual change in the unemployment rate. Both coefficients are highly significant and are consistent with those obtained by Okun's original work with quarterly data and more recent work using annual data (Knotek 2007, 80-81).

The estimated regression equation implies that a year of zero GDP growth would cause the unemployment rate to rise a little over 1.3 points and a one percentage point increase in the output growth rate would be associated with an unemployment rate that would be almost 0.4 points lower than otherwise. These estimates are very close to the 1.2 and −0.35 estimates obtained by Knotek (2007) using annual data from the 1949-2006 period, and both sets of estimates suggest that the growth rate of real GDP that is consistent with a steady unemployment rate is about 3.4 percent.[3]

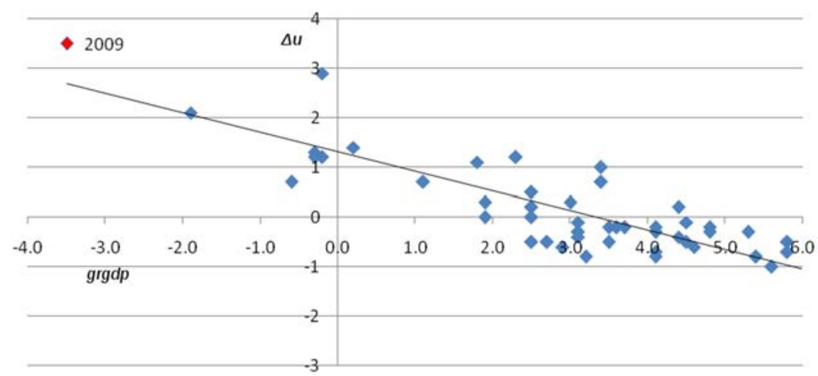
R-squared measures the proportion of the variation in annual changes in U that can be explained by variation in real GDP growth over the last 50 years. Since almost three-quarters of observed changes in U can be explained by fluctuations in

3. Since $\Delta u = \alpha + \beta * grgdp$, the output growth rate that keeps U constant = $-\alpha/\beta$. My estimate of this growth rate is 3.36 percent per year while Knotek's estimate obtained with data from an earlier period was 3.43.

GDP growth rates, it is clear that this strong cyclical variation in the unemployment rate must be taken into account in any attempt to measure how much of the current relatively high unemployment rate is due to the way job seekers have responded to an extension of UI benefits.

These regression results can be used to control for the impact of the business cycle on unemployment in coming to a rough estimate of the potential impact of extended UI, ceteris paribus. Figure 1 shows the scatter plot of the 50 annual observations[4] and the resulting regression line. Any point on the downward-sloping regression line in Figure 1 shows the annual change in the unemployment rate that one would expect to observe given a growth rate in output.

Figure 1. Δu = f(*grgdp*) 1961-2010 annual data



The point labeled 2009 in the upper left corner of the scatter diagram shows the unemployment rate change-GDP growth combination for 2009—the worse year of the recession—when real GDP fell by 3.5 percent. Such a decline in GDP would be expected to raise U by 2.7 points above the prior year's level. However, the actual increase in unemployment from 2008 was 3.5 points. This 0.8 point excess of the observed increase in U over what one would expect given the magnitude of the decline in real GDP could be due to the significant increase in the maximum duration of UI from 13 weeks up to 53 weeks that took place in 2009 (U.S. Department of Labor 2011). This rough estimate of the increase in U that could be due to extended UI is at the bottom of the 0.8 to 1.2 point range found by Mazumder and others but it is less than one-third of the 2.7 point increase suggested by Barro (2010).

4. One year's observation is lost as the unemployment rate variable is the change in U between consecutive years.

An alternative measure of the severity of the 2008-09 recession and its impact on the unemployment rate was obtained by applying the Hodrick-Prescott filter (Hodrick and Prescott 1997) to the GDP (Backus and Kehoe 1992) and unemployment rate data (Mocan 1999) to obtain their HP-filtered trends.[5] The cyclical components of GDP and U were then calculated by subtracting the annual observed values from trend. The cyclical component of unemployment (*ugap*) was regressed on GDP's deviation from trend (*gdpgap*) expressed as a percentage of its trend value. The regression results are shown in Equation 2 with the T ratios in parentheses. The R-squared for the estimated equation was .61.

$$\text{Equation 2: } ugap = \underset{(-.03)}{-0.003} - \underset{(-8.84)}{0.508}{*}gdpgap$$

The slope coefficient on *gdpgap* provides an estimate of how much the unemployment rate deviates from trend for each one percentage point that output deviates from trend. The statistically insignificant intercept term is consistent with the expectation that unemployment will not deviate from its natural level if real GDP does not deviate from trend.

Figure 2 shows the scatter plot for the 1960-2010 annual *ugap* and *gdpgap* observations along with the linear function representing the estimated regression equation. The data point for 2009 is identified, and its deviation from the estimated regression line suggests that the unemployment rate was higher than would be expected given the size of the output gap in than year. Cyclical unemployment is estimated to be 2.0 percentage points when 2009 real GDP is 3.9 points below trend, but the observed gap between the actual unemployment rate and trend was 3 points. This one point excess of the observed *ugap* over what would be expected provides another rough estimate of how much higher unemployment could be a consequence of extended UI. This estimate is also consistent with those found by Mazumder but much less than the 2.7 point increase suggested by Barro.

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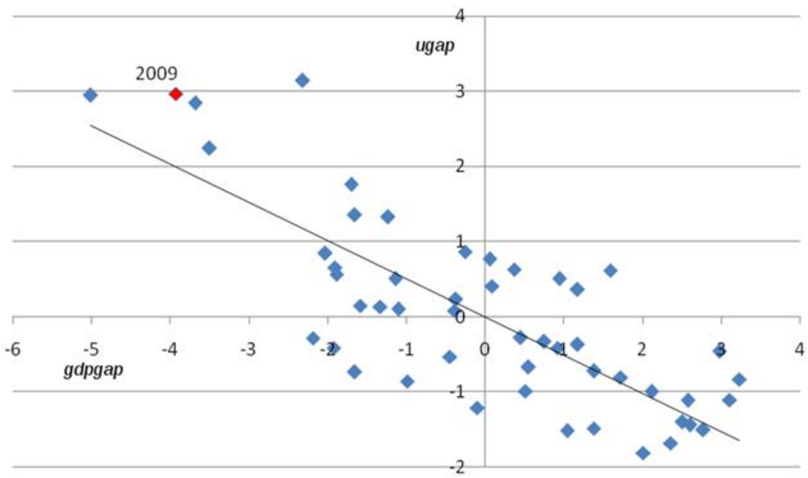
5. The trends in U and GDP identified by the HP-filter are heavily dependent upon the weight applied to the long-run growth component in the data series. This weight, denoted in the literature as lambda, varies with the frequency of observations. Following Mocan's work on cyclical unemployment and Bacus' work on GDP, lambda was set equal to 1600 for the unemployment series and 100 for the GDP series. The debate over appropriate values for this weight is beyond the scope of this paper. See Ravn and Uhlig (2002) for a discussion of the issues.

Figure 2. *ugap* = f(*gdpgap*) 1960-2010 annual data



A critical limitation of these simple regression models is that there are a number of factors—in addition to the extension of UI—not included in the model specified by Equations 1 and 2 that could have an impact on the unemployment rate. Consequently, the models were estimated again this time adding an additional explanatory variable in an attempt to measure more precisely the impact of the extension of UI benefits on the unemployment rate while controlling for GDP.

Data obtained from the U.S. Department of Labor documenting the effective date of special UI benefit extensions and the length of those extensions was used to construct a variable (*xui*) measuring the number of additional weeks of unemployment benefits that were available to covered workers during each year over the 1960-2010 period (U.S. Department Of Labor 2011). During some years, extended UI benefits were not available over the entire year, in which case the number of extended weeks of benefits was calculated by multiplying the maximum number of weeks that benefits were payable by the fraction of the year they were available.[6]

The simple inclusion of the annual change in *xui* (Δ*xui*) in Equation 1 or *xui* in Equation 2 as additional explanatory variables is likely to produce biased

6. The *xui* variable measures the number of weeks of extended unemployment benefits that were available over the year. In years when no extended benefits were available *xui* was set to 0. During the two years following the 9-11 attacks, for example, UI was extended by 26 weeks beginning in March of 2002 and those extensions were continued until the end of 2003 at which point in time they were phased out. The 2002 value of *xui* was set equal to 10/12*26 = 21.67 since 26 weeks of benefits were only available for 10 months of that year. The 2003 value for *xui* was set equal to 26 because 26 weeks of extended benefits were available for that entire year. Each of the other annual values for *xui* was established in the same manner.

estimates of the impact of changing durations of UI benefits on the unemployment rate because annual changes in *xui* are sure to have both exogenous and endogenous components. OLS regression is expected to provide unbiased estimates of the impact of extended UI on the unemployment rate under the assumption that *xui* is exogenous and the direction of causation runs only from changes in weeks of extended UI benefits to changes in the unemployment rate. A correct specification of the model must take into account the endogenous behavior of the duration of unemployment benefits because legislation extending UI benefits is clearly influenced by the state of the economy and the actual unemployment rate, although with a consistent lag (Coven and Stone 2008).

Two versions of each of the models defined by Equations 1 and 2 are specified with the addition of an extended UI independent variable in ways that avoid this likely bias. The first version of this specification—Okun1 and HP1—include Δxui and *xui* as additional independent variables respectively, but also include the lagged value of the dependent unemployment variable as a third independent variable in order to control for the impact of previous labor market conditions that could inflate the estimate of the impact of UI on U. The resulting specifications of the two models, which include both a UI and a lagged dependent variable on the right hand side, are:

Okun1: $\Delta u = \alpha + \beta * grgdp + \gamma * \Delta xui + \delta * lag\, \Delta u$

HP1: $ugap = \alpha + \beta * gdpgap + \gamma * xui + \delta * lag\ ugap$

A second version of the two models adds a UI variable that is constructed to remove the endogenous component of variations in UI policy so that the new variable captures only exogenous variation in the extent of UI benefits. This variable was obtained by regressing *xui* (weeks of extended UI eligibility) on lagged values of the GDP gap and U gap variables and saving the residuals as a new variable, *xuishock*. This new variable should better capture variation in the extent of UI benefits that are due to changes in the national political landscape over time.[7] This new variable is a 'generated regressor,' and Pagan (1984) demonstrated that OLS regression models containing explanatory variables derived from predicted and residual values from prior regressions suffer from a version of the errors-in-variables problem. Consequently, the standard errors and resulting t ratios should be used with caution.[8]

The addition of *xuishock* and Δ*xuishock*—the annual change in *xuishock*—results in the following specifications of the two models.

7. As a test of this expectation, *xuishock* was regressed on three dummy variables representing Democratic Party control of the Presidency, the House and the Senate. Democratic control over the Senate was found to be positively and significantly related to *xuishock*, but the F test did not indicate joint significance.

Okun2: $\Delta u = \alpha + \beta * grgdp + \gamma * \Delta xuishock$
HP2: $ugap = \alpha + \beta * gdpgap + \gamma * xuishock$

Of particular interest for this research is the sign and significance of γ, the slope coefficient on the variables measuring the weeks of extended UI eligibility. In the two Okun versions of the model, the γ coefficient denotes how much the unemployment rate would change in a given year for every one additional week that UI benefits are extended, ceteris paribus. The γ term in the two HP versions express a similar relationship, only in terms of unemployment's deviation from trend for every week of extended UI. Each of the models was estimated with OLS regression and the results are presented in Table 1 below.

TABLE 1. Summary Regression Results (t stats)

Okun Model Dependent Variable = Change in Unemployment Rate
HP Model Dependent Variable = Unemployment Rate Deviation from HP Trend

Model	GDP	UI	lagU	Adj. R^2	F	D-W
Okun1	-0.334 (-11.60)	0.027 (4.25)	.044 (0.68)	0.84	82.4	NA
Okun2	-0.346 (-10.91)	0.025 (3.44)	NA	0.80	95.4	1.69
HP1	-0.337 (-7.53)	0.029 (3.87)	0.36 (4.94)	0.83	81.8	NA
HP2	-0.477 (-8.48)	0.025 (2.38)	NA	0.66	48.9	0.40
HP2GLS	-0.391 (-10.10)	0.029 (3.95)	NA	0.78	87.0	1.66

The estimated impact of extended UI benefit durations is very consistent in all four models, and slope coefficient estimates on the GDP and UI variables are all of the expected signs and are statistically significant at the alpha=.011 level or better. The estimates of γ obtained from the Okun specification of the model suggest that each one week extension of UI benefits raises the unemployment rate by 0.025 to .027 points, ceteris paribus. The HP specifications are quite consistent as well, with each additional week of extended UI estimated to raise unemployment's deviation from trend by .025 to .029 points.

Since time series regression analysis is typically prone to serial correlation resulting in untrustworthy standard errors and t-ratios, the residuals were examined and the Durbin-Watson test statistic for autocorrelation was calculated for the Okun2 and HP2 models. The Durbin-Watson *d* is not reported for the Okun1 and HP1 models because it is not a valid test for autocorrelation when the model contains the lagged dependent variable as a regressor (Durbin 1970), so an

8. There are additional technical issues associated with the use of generated regressors that go beyond the scope of this paper. My thanks go to a referee for bringing them to my attention. Interested readers should see Oxley and McAleer (1993) for details.

alternative test was used for those models.[9] The Okun versions of the model demonstrated no evidence of serial correlation—most likely due to the first-differencing of the data in those models—but this was not the case for the models using the HP filtered data. Durbin's alternative test confirmed the presence of autocorrelation in the HP1 model and examination of the residuals from the HP2 regression combined with a *d* statistic much lower than the critical value for d_l confirmed the presence of first-order autocorrelation in that model as well (Gujarati 2006; Studenmund 2006). In order to correct for serial correlation, HP2 was re-estimated using generalized least squares with the following specification.

HP2GLS:

$$ugap_t - \varrho^* ugap_{t-1} = \alpha + \beta^*(gdpgap_t - \varrho^* gdpgap_{t-1}) + \gamma^*(xuishock_t - \varrho^* xuishock_{t-1})$$

Rho (ϱ) was estimated to be .80 and the resulting GLS regression results are provided in the last row of Table 1.[10] The residuals from this regression showed no sign of first-order autocorrelation, and *d* increased to 1.66, which exceeds d_u, the upper bound of the indeterminate range for *d*. The resulting estimates of the impact of the GDP gap and extended weeks of UI on the unemployment rate are still significant and of the expected signs.

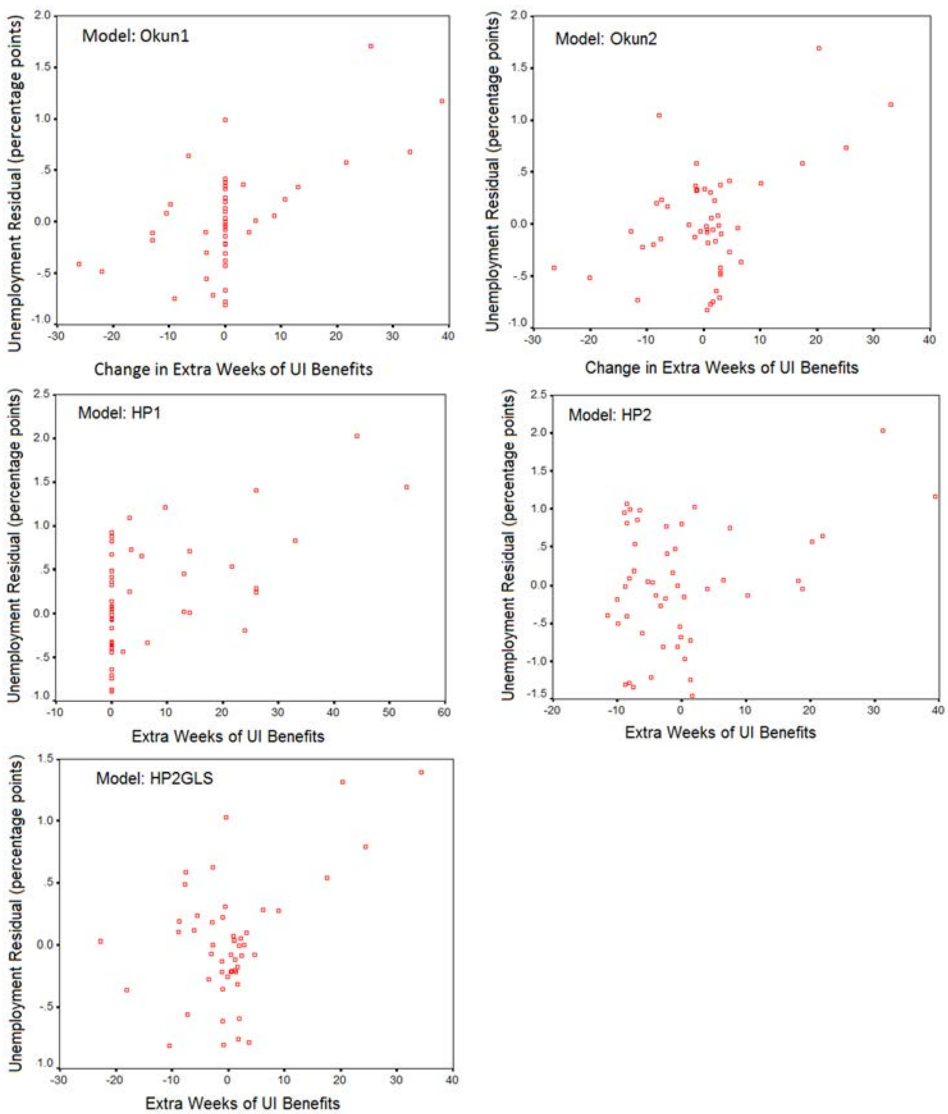
A visual indication of the relationship between extended UI benefits and unemployment is provided by the partial residual plots in Figure 3 (Larsen and McCleary 1972). Each of the five plots—one for every model specification—is formed as $res + \gamma^* ui$ where *res* are the regression residuals from the model and γ is the regression coefficient on *ui*, the relevant variable from each model.

These regressions attempt to measure the impact of extended UI on the unemployment rate holding GDP constant. If it were the case that extended benefits reduced GDP while they raised unemployment, my estimate of the impact of UI on unemployment would be biased downward. But a body of research suggests that this scenario is unlikely, as the majority of the unemployed whose UI benefits are exhausted leave the labor force instead of going back to work (Card et al. 2007; Rothstein 2011). Those individuals whose employment is not influenced by extra UI are certainly picked up in my regressions.

9. The residuals from the Okun1 and HP1 OLS regressions were regressed on all regressors in each model plus the lag of the residuals. The t ratio for the coefficient on the lagged residual was then used to test for the presence of serial correlation.

10. An estimate of ϱ was obtained by regressing the residuals from the HP2 regression on their lagged values and from the Durbin-Watson test statistic. Both methods resulted in the same value.

Figure 3. Partial residual plots for the UI variable



How Do These Results Compare to Barro's Estimate of the Impact of Extended UI?

The regression results described in the previous section were used to calculate how much higher the unemployment rate was in 2010 as a consequence of the extension of UI benefits, ceteris paribus. Table 2 shows the results of these calculations. The regression coefficient on the UI variable in each of the five specifications of the model was multiplied by the relevant independent UI variable to estimate how much the unemployment rate was increased as a consequence of the increases in the duration of UI benefits that occurred since the Emergency Unemployment Compensation program was created in the summer of 2008. The estimates derived from the two Okun models are percentage point increases in the unemployment rate in 2010, while the HP model estimates are the percentage point increases in the excess of the observed unemployment rate over the trend unemployment rate. These estimates are larger than those provided by Rothstein (2011) but are very close to the 0.8-1.7 point range obtained by Mazumder (2011, 2). They are, however, lower than the 2.7 point estimate provided by Barro (2010).

TABLE 2. Estimated Impact on 2010 Unemployment Rate of Extended UI

Model	Increase in U
Okun 1	1.43
Okun 2	0.77
HP 1	1.54
HP 2	0.78
HP2GLS	0.90

Summary and Conclusions

The models developed in the previous section estimate that the unemployment rate would have been between 0.77 and 1.54 points lower in 2010 had it not been for the unprecedented increases in the duration of UI benefits that occurred since the Emergency Unemployment Compensation program was created in the summer of 2008. These results are contrasted with those provided by Barro (2010), who assumed that all of the observed increase in long-term unemployment during this recession was due to extended UI, and that if it were not for those extensions the unemployment rate would have been 6.8 percent in the summer of 2010 instead of 9.5. Another estimate (Mazumder 2011) based on steady-state unemployment theory found that the unemployment rate would have been in the 7.8 to 8.7 percent range in the absence of extended UI benefit durations. The present paper controls for the severity of the recession and finds that the 2010 unemployment rate would have been between 8.0 and 8.8 percent if it were not for extended UI benefits.

While my research indicates that Barro's assessment of the impact of UI on unemployment may be exaggerated, one important point to glean from all of these estimates is that policymakers have good reasons to use caution when considering the unemployment rate as a guide to macroeconomic policy, especially during periods when an extension of UI benefits has pushed the unemployment rate above levels that would have prevailed otherwise.

Appendices: Data and Regression Output Files

GDP and U detrended with HP Filter This Excel file contains real GDP and the unemployment rate for the 1960-2010 period. Deviations in real GDP from trend (as a % of trend GDP) and deviations in U from its trend are calculated using the Hodrick-Prescott Filter.

SPSS file with dependent and independent variables for all regressions This file contains all of the dependent and independent variables used in the Okun and HP models as well as the transformed data for the GLS estimate of the HP2 model. The residuals from all models are also saved here and used to create the partial residual plots.

These files contain the SPSS output for each of the five regression models.

SPSS Output for Okun 1 Model
SPSS Output for Okun 2 Model
SPSS Output for HP 1 Model
SPSS Output for HP 2 Model
SPSS Output for HP 2 GLS Model

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