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The Wage Curve Reloaded

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ABSTRACT

The Wage Curve Reloaded*

This paper provides evidence for the existence of a wage curve – a micro-econometric association between the level of pay and the local unemployment rate – in modern U.S. data. Consistent with recent evidence from more than 40 other countries, the wage curve in the United States has a long-run elasticity of approximately -0.1. In line with the paper's theoretical framework: (i) wages are higher in states with more generous unemployment benefits, (ii) the perceived probability of job-finding is lower in states with higher unemployment, and (iii) employees are less happy in states that have higher unemployment. We conclude that it is reasonable to view the wage curve as an empirical law of economics.

JEL Classification: J3, E2

Keywords: wages, unemployment, wage curves

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The Wage Curve Reloaded

1. Introduction

One of the oldest empirical questions in economics is that of how the price of labor is affected by the unemployment rate. Following a tradition begun by the New Zealand economist A.W. Phillips (1958), this question has traditionally been studied with aggregate time-series methods. Although its robustness is still debated, the Phillips Curve, which is usually conceived of as a relationship between wage growth and unemployment, has become part of standard macroeconomics.

Blanchflower and Oswald (in, for example, 1994a) argue instead for the use of microeconomic data.[1] Their book documents the existence of a logarithmic curve linking the level of pay to the unemployment rate in the local area. Its conclusion is that, in sixteen nations, including the United States, the data are well described by a wage curve with an unemployment elasticity of approximately –0.1.[2] A doubling of the unemployment rate is then associated with a ten percent decline in the level of the (real) wage. This is presented in the book as an empirical law of economics.

Since then, Blanchflower and Oswald's conclusions have been scrutinized by other researchers. The wage-curve finding has now been replicated for more than 40 nations. Its existence in the United States, however, is still viewed as controversial.

One reason is that Blanchard and Katz (1997) argue for a Phillips Curve, rather than a wage curve, in United States data. Staiger, Stock and Watson (2002) and Card and Hyslop (1997) also report a high level of auto-regression in U.S. wages. In contrast, Hines, Hoynes and Krueger (2001) conclude that a wage curve specification has a more natural theoretical interpretation and fits the data (hours as well as wages) better than the Phillips Curve specification. Hines et al (2001) produce evidence of wage curves using annual and hourly earnings from the 1977-2000 March CPS files. The authors also uncover wage curves in the PSID for male household heads aged 25-59 who were continuously employed for 1967-1987 and 1977-1996. Using the PSID, Hines et al suggest that a 3 percentage point decline in the unemployment rate is associated with a 4 per cent increase in real wages, which translates into an elasticity similar to the Blanchflower-Oswald number.

In principle, the wage curve should be as interesting to macroeconomists as to labor economists. Yet the empirical results on the wage curve's existence have so far made relatively little impression on American macroeconomics. This does not seem ideal. First, macroeconomic analysis has for some decades stressed the need for microeconomic foundations. It might be

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[1] See also Blanchflower and Oswald (1990, 1993, 1994b, 1995, 1996, 2000).

[2] In Blanchflower and Oswald (1994a) we used micro-data for twelve countries: the United States, the United Kingdom, West Germany, Austria, the Netherlands, Switzerland, South Korea, Norway, Ireland, Italy, Canada and Australia. The book also discussed evidence based on other researchers' work in a further four countries - Sweden, the Côte d'Ivoire, Japan and India.

natural to expect such strictures to apply empirically as well as theoretically. Second, some macroeconomics textbooks, including Blanchard (2003), make extensive theoretical use of a wage curve (at the aggregate level, which he refers to as a 'wage-setting' relation), but do not lay out the evidence for such a curve.[3] Third, the idea of viewing different US states as usefully different mini laboratories for study -- effectively what chapter 4 of Blanchflower and Oswald (1994a) proposes -- has become conventional in various areas of economics.

Reconsidering the evidence

The literature on international wage curves is now so large that the case of the United States is becoming increasingly anomalous. It needs to be reconsidered.

Subsequent to the publication of the Blanchflower-Oswald book [4], wage curves were reported for Argentina, Australia, Austria, Belarus, Belgium, Brazil, Bulgaria, Burkina Faso, Canada, Chile, China, Côte d'Ivoire, Czech Republic, Denmark, East Germany, Estonia, Finland, France, Great Britain/UK, Holland, Hungary, India, Ireland, Italy, Japan, Latvia, New Zealand, Norway, Poland, Portugal, Romania, Russia, Slovakia, Slovenia, South Africa, South Korea, Spain, Sweden, Switzerland, Taiwan, Turkey, USA, and West Germany. The list below provides a summary.

Wage Curves in the Research Literature

1. *Argentina*	Galiani (1999).
2. *Australia*	Blanchflower and Oswald (1994a); Kennedy and Borland (2000).
3. *Austria*	Blanchflower and Oswald (1994a) (2000); Winter-Ebmer (1996).
4. *Belarus*	Pastore and Verashchagina (2004).
5. *Belgium*	Janssens and Konings (1998).
6. *Brazil*	Amadeo and Camargo (1997); Barros and Mendonca (1994).
7. *Bulgaria*	Blanchflower (2001); Iara and Traistaru (2004).
8. *Burkina Faso*	Lachaud (1998).
9. *Canada*	Blanchflower and Oswald (1994a) (2000); Fares (2002); Leung (1997).
10. *Chile*	Berg and Contreras (2004).
11. *China*	Wu (2004); Sabin (1999).
12. *Côte d'Ivoire*	Hoddinott (1996).
13. *Czech Republic*	Blanchflower (2001); Huitfeldt (2001); Galušcák and Münich (2002).
14. *Denmark*	Nicolaisen and Tranaes (1996).
15. *East Germany*	Blanchflower (2001); Baltagi, Blien and Wolf (2000); Buscher (2003); Elhorst, Blien and Wolf (2002); Pannenberg and Schwarz (1996).
16. *Estonia*	Blanchflower (2001).

[3] Blanchard (2003), pp. 126-128 and 273-275.

[4] There have also been various theoretical contributions, including Campbell and Orszag (1998), Johansen (1997), Knight (1998), Ma (2004), Mankiw and Reis (2003), Roberts (1997), Sato (2000), Sessions (1993) and Whelan (1997). Some papers, such as Guichard and Laffargue (2000), now use the term 'wage curve' when looking at the macroeconomic effects of unemployment within a panel of countries; but we try here to reserve the term for micro-econometric and disaggregated estimation.

17. *Finland*	Pekkarinen (2001).
18. *France*	Montuenga et al (2003); Estevao and Nargis (2001); Gianella (2002); Glaude and L'Héritier (1993), Delteil et al (2004).
19. *Great Britain/ United Kingdom*	Blanchflower and Oswald (1990) (1994a) (1994b) (2000); Montuenga et al (2003); Bell, Nickell and Quintini (2002); Blackaby and Hunt (1992); Black and Fitzroy (2000); Barth et al (2002a, b); Collier (2001).
20. *Hungary*	Blanchflower (2001); Kollo (1998); Kertesi and Kollo (1997) (1999); Delteil et al (2004).
21. *India*	Bhalotra (1993).
22. *Ireland*	Blanchflower and Oswald (1994a).
23. *Italy*	Blanchflower and Oswald (1994a) (2000); Montuenga et al (2003); Chiarini and Piselli (1997); Canziani (1997).
24. *Japan*	Rebick (1993); Montgomery (1994); Kano (2003).
25. *Latvia*	Blanchflower (2001); Adaimate (2001).
26. *Netherlands*	Blanchflower and Oswald (1994a) (2000); Groot et al (1992); Graafland (1992).
27. *New Zealand*	Papps (2001); Morrisson and Poot (1999).
28. *Norway*	Blanchflower and Oswald (1994a) (2000); Johansen (1995) (1999); Falch (2001); Bardsen et al (2004); Brunstad and Dyrstad (1997); Dyrstad and Johanssen (2000); Wulfsberg (1997); Johansen et al (2001); Barth et al (2002a, b).
29. *Poland*	Blanchflower (2001); Iara and Traistaru (2004); Duffy and Walsh (2000) (2002); Basu et al (1995).
30. *Portugal*	Montuenga et al (2003).
31. *Romania*	Kallai and Traistaru (2001).
32. *Russia*	Blanchflower (2001).
33. *Slovakia*	Blanchflower (2001); Basu et al (1995) (2000); Huitfeldt (2001).
34. *Slovenia*	Simoncic and Pfajfar (2004).
35. *South Africa*	Kingdon and Knight (1998).
36. *South Korea*	Blanchflower and Oswald (1994a).
37. *Spain*	Garcia-Mainar and Montuenga-Gomez (2003); Jimeno and Bentolila (1998); Canziani (1997); Sanromá and Ramos (2003); Bajo et al (1999).
38. *Sweden*	Holmlund and Skedinger (1990); Edin and Holmlund (1989); Edin, Holmlund and Ostros (1994).
39. *Switzerland*	Blanchflower and Oswald (1994a) (2000).
40. *Taiwan*	Rodgers and Nataraj (1999).
41. *Turkey*	Ilkkaracan and Raziye (2003).
42. *USA*	Blanchflower and Oswald (1990) (1993) (1994a) (1995) (1996) (2000); Blanchflower, Oswald and Sanfey (1996); Blanchard and Katz (1997); Card (1995); Bratsberg and Turunen (1996); Deller and Tsai (1998); Deller and Zhing (2003); Boushey (2002); Turunen (1998); Barth et al (2002a, b); Bertrand (2004); Hines, Hoynes and Krueger (2001); Bartik (2000).
43. *West Germany*	Blanchflower and Oswald (1994a) (1996); Bellmann and Blien (2001); Wagner (1994); Baltagi and Blien (1998); Buettner (1999); Longhi et al

(2004); Pannenberg and Schwarze (1998); Blien (2003); Rendtel and Schwarze (1996); Baltagi et al (2000).

What can be learned from this work? Although inevitably imperfect, the statistical law of a wage curve with an elasticity of –0.1 has turned out to fit the data surprisingly well.[5] Interestingly, results for the Scandinavian countries emerge somewhat below the -0.1 estimate (e.g. Pekkarinen (2001); Brunstad and Dyrstad (1997); Holmlund and Skedinger (1990)). There is also evidence from one or two of the former communist countries of an elasticity far above the -0.1 estimate (e.g. Blanchflower (2001); Pastore and Verashchagina (2004); and Pannenberg and Schwarz (1998)) although other papers for these countries do find numbers close to -0.1 (such as Kertesi and Kollo (1997, 1999); Duffy and Walsh (2002); Elhorst et al. (2002); and Kallai and Traistaru (2001)).

A recent meta-analysis -- on a sample of 208 wage/unemployment wage curve elasticities from the literature -- by Nijkamp and Poot (2005) concludes that

> "the wage curve is a robust empirical phenomenon … but there is … evidence of publication bias. There is indeed an uncorrected mean estimate of about -0.1 for the elasticity. After controlling for publication bias by means of two different methods, we estimate that the 'true' wage curve elasticity at the means of study characteristics is about -0.07".

Nijkamp and Poot (2005) also test for, and find no evidence of, what they term a Blanchflower/Oswald 'advocacy' effect.

Most economists are unlikely to feel strongly about the possible difference between a wage curve elasticity estimate of -0.07 and one of -0.1. What matters more is whether there are countries in which a wage curve cannot reliably be found. Here the case of the United States needs to be confronted.

2. The Data Reloaded

Is there, or is there not, a U.S. wage curve? Given that wage curves have been found for many dozens of other nations, it would be remarkable, and presumably important, if the same were not true in the United States.

The aim of this paper, therefore, is to study the patterns in modern American data. We re-estimate the U.S. wage curve on a state level panel using CPS data from 1980-2001. We explore points made by Blanchard and Katz (1997).

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[5] On page 357 of Blanchflower and Oswald (1994a) it is argued that:

> "Future work will have to begin to test for statistically significant differences among numbers that lie in a rough band from -0.05 to -0.20. It would probably be unwise to treat the minus-point-one rule as more than one of thumb".

The paper tests explicitly for the effect of unemployment upon wages and does not discuss the possible structure of a classical labor-supply function. This is because: first, the paper takes the efficiency-wage theory as its maintained hypothesis and in such a framework a no-shirking condition replaces the labor supply curve; second, the paper aims to be consistent with the literature on Phillips Curves and wage curves; third, in other results not reported here we have found that unemployment variables enter wage equations more strongly than do employment-rate variables.

Overall, our results suggest that the United States has more autoregression in pay than many nations, but that there is much evidence of a U.S. wage curve. This has implications that seem to matter. It means, first, that US labor markets may not be fundamentally different from other nations' labor markets (contrary to what has sometimes been claimed), and, second, that American macroeconomists who use the Phillips Curve in modeling might wish to work more with micro data.

With the advantage of hindsight, our 1994 book failed to examine sufficiently carefully the autoregressive nature of hourly pay in the United States.

Any labor economist would argue that wages and unemployment are simultaneously determined. Within the wage-curve debate, therefore, issues of identification and measurement will matter. Moreover, given that wages are famously sticky, it is unlikely that it will be straightforward to distinguish between short-run and long-run effects on pay. Wage dynamics can be expected to play a central role.

Put informally, four views are visible in the economics literature.

Idea 1 (Phillips). The rate of growth of pay is inversely related to the unemployment rate. This hypothesis is explored in Phillips' 1958 Economica paper. He used a long run of aggregate time series data for Britain. The idea of a function dw/dt = f(U) is not necessarily a-theoretical. It can be justified by appealing to the notion that excess supply of a commodity gradually pushes down the price of that commodity. Here f(.) has a negative gradient. But Phillips was not clear about whether any relationship was to be expected between the level of wages, w, and unemployment, U. Sargan (1964) was later to argue a version of that.

Idea 2 (Harris-Todaro) To take a job in a region or industry (or in principle even a country) that experiences a lot of unemployment, the typical worker needs to be paid well. This is an application of the concept of compensating differentials. A state like West Virginia, which traditionally has had one of the highest rates of joblessness in the United States, should in equilibrium offer better pay than one with persistently low unemployment. Otherwise workers will find life too risky to stay. At the region or industry level, therefore, w = g(U). Here the function g(.) will have a positive gradient.

Idea 3 (The wage curve). In most countries, it has been found, following Blanchflower and Oswald, that regions with higher unemployment have lower wages. One way to rationalize such a discovery is to appeal to non-competitive theories of the labor market – for example to the idea of a no-shirking condition or a bargaining-power effect. According to this kind of analytical framework, higher local unemployment makes life tougher for workers (for example, they will find it harder to obtain work if laid off by their current employer), and therefore it is not necessary for employers to remunerate those individuals so well. Here w = h(U), and the h(.) function has a negative gradient.

Idea 4 (Labor demand and supply). If unemployment can in some circumstances be viewed as the negative of employment, an observed function w = j(U) might be construed as a labor demand or supply function. This logic might be thought to be a stretch, because labor demand and supply theory is not typically framed in unemployment space. Nevertheless, a function w = j(U) might be seen as some kind of demand or supply curve written in mild disguise. If j(U) is a miss-specified labor supply function, for example, it might be argued that j(U) would have a negative gradient. If it is a miss-specified demand function, it might be expected to have a positive gradient.

Among these four views, there is currently disagreement, and perhaps even confusion, in labor economics and macroeconomics. We think the current position is this. The average US labor economist believes in Harris-Todaro and, with a bit of skepticism along the way about the nature of macroeconomic data, also believes in some form of Phillips Curve. The average European labour economist does not believe in Harris-Todaro and, with a bit more respect along the way for macroeconomic data, believes in some form of wage curve.

Recent debate has probably been hampered by a tendency for researchers to concentrate on extremes. Some commentators, for example, interpret Blanchflower and Oswald (1994a) as arguing for a necessarily static wage curve without any autoregression in wages (although the book includes estimated equations with lags on the dependent variable). Similarly, because Phillips's original article does not discuss the possibility of dynamics other than a unit root in the pay equation, the macroeconomics literature has been dominated by simple functional forms.

We aim to find common ground between disagreements in Blanchard and Katz (1997, 1999), Staiger, Stock and Watson (2002), Card and Hyslop (1997), Blanchflower and Oswald (1994a), Bell, Nickell and Quintini (2002), Barth et al (2002a, b), and Fares (2002).

3. A theoretical framework

If wages and unemployment are simultaneously determined, how should we think about the spatial patterns in pay and joblessness? A model favored by Blanchflower and Oswald (1994a) draws on efficiency wage theory. An illustrative version is as follows.

Consider an economy consisting of just two regions. The following assumptions are made about region 1 and, with small modifications, also about region 2.

A1 Assume that workers are risk-neutral. They get utility from income and disutility from effort. Define the wage as w and the level of on-the-job effort as e. Assume that a worker's utility equals the simple difference between income and effort, so that utility is u = w – e.

A2 Assume that effort at work, e, is a fixed number determined by technology, but that individual employees can decide to "shirk" and exert zero effort. If undetected by the firm, these individuals earn wage w and have e = 0, so that u = w. They are then better off than employees who provide effort.

A3 An individual who shirks runs the risk of being detected. Designate as δ the probability of successfully shirking, that is, of escaping detection. Assume that anyone caught shirking is fired, and has then to find work elsewhere (at required effort e). Let the expected utility of a fired worker be $\overline{w}$. Define it

$$\overline{w} = (w - e)\alpha(U) + b[1 - \alpha(U)]. \qquad (1)$$

This is a convex combination of w – e, the utility from working at the required effort level, and of b, which is defined as the income value of unemployment benefit plus leisure. The function α(U) measures the probability of finding work, and how that is affected by the level of unemployment, U, prevailing in the local labor market.

A4 Assume that there is a constant rate of break-up, r, of firms. In steady-state equilibrium, total new hires in the local economy are α[L – n] where L is working population and n is employment, and rn = α[L – n].

Unemployment is by definition $U \equiv 1 - \frac{n}{L}$, so $r = \frac{r}{U} - \alpha$

This implicitly defines a function α(U) where the first derivative is negative and the second derivative is positive. Thus the probability of finding a job, α, is a convex function of the unemployment rate, U.

A5 Equivalent equations hold in the second region.

A6 However, the second region differs from the first in that both workers and non-workers enjoy a non-pecuniary benefit, ø, from living in the region.

A7 Each region is affected by shocks to the demand for labor. The shock variable is denoted s in region 1, with a density function g(s). There is an equivalent shock variable in region 2, with density h(.). These shocks could be thought of as exogenous movements in real input prices, but other interpretations are possible.

A8 Workers are free, between periods, to choose to live in whichever region they prefer. They cannot migrate during a period.

The assumptions given above describe a form of efficiency-wage model. The model's key characteristic is that employers must pay a wage that is sufficiently high to induce employees not to shirk. In equilibrium, workers must be behaving optimally in their effort decisions, and firms must be behaving optimally in their wage-setting. Regions differ in their non-pecuniary attractions: one of the two is a nicer place to live than the other. Excluding degenerate equilibria, however, each region must offer workers the same level of expected utility. This condition defines a zero-migration equilibrium.

Blanchflower and Oswald (1994a) show that the following result can be proved.

Proposition 1: *Each region has a downward-sloping convex wage curve. If regions have the same level of unemployment benefit, they have a common wage curve given by the equation*

$$w = b + e + \frac{e\delta}{(1-\delta)[1-\alpha(U)]} \qquad (2)$$

The convexity of the wage curve follows from the convexity of the $\alpha(U)$ function, and can be checked by differentiation. Intuitively, as unemployment U rises, firms realize that their employees are inherently more frightened of losing their jobs, and can pay lower levels of remuneration while maintaining the necessary degree of worker effort. A corollary is:

Proposition 2: *There is involuntary unemployment in equilibrium.*

The details of the proofs, and other associated corollaries, are given on pages 66-69 of Blanchflower and Oswald (1994a).

4. Empirical Issues

In this paper we attempt to implement empirically a version of Equation 2.

First, a logarithmic wage equation is estimated. This is not essential but follows convention in the labor literature. Second, we test for the influence of unemployment benefits, b, within a wage equation. Although it follows directly from the theory, we have been unable to find a literature on testing for the effects of benefit levels in regional or state-level wage equations. The previous wage-curve literature has been unable to do so, but in US data it is possible to use information on state unemployment insurance payments to construct a value for the b variable across space. If information on b is not available, it might be argued, as in Blanchard and Katz (1999), that the coefficient on any lagged dependent variable will be biased. Third, following the previous wage-curve literature, the nonlinear function embedded within equation 2 is approximated empirically by a logarithmic function of unemployment, U. This is primarily for simplicity. Fourth, equations are estimated with a lagged dependent variable (that is, a regressor w_{t-1}). This is to measure the degree of autoregression in pay, and to allow a direct check for pure Phillips Curve specifications. Fifth, implicit in the analysis will be the assumption that joblessness makes people unhappy, that is, that unemployment benefits b do not make up sufficiently for the drop in utility upon loss of a job; this is consistent with empirical evidence on psychological wellbeing, such as that in Clark and Oswald (1994).

Identification issues have to be addressed. Here the form of wage equation makes it natural to use benefits b as an instrument for wages w. To complete the model, an unemployment equation is also needed. We assume it is a second-order AR process, derived from a neoclassical demand curve for labor, and use two lags on U as instruments for unemployment. Later in the paper we estimate this function $U_t = \Omega(U_{t-1}, U_{t-2}, w_t, \text{controls})$ directly.

The paper begins with simple wage equations. They are estimated using state-year weighted averages based on the Outgoing Rotation Group (ORG) files of the Current Population Survey. The data cover the years 1979-2001. The dependent variable is the log of hourly earnings in each state-year cell. Hourly earnings are calculated as earnings per week divided by usual hours. Top-coded earnings are multiplied by 1.5. Weights are used to calculate state-year cell averages for the independent and dependent variables.

There is a potential difficulty with the ORG earnings or wage data. Hirsch and Schumacher (2004) have pointed out that over 30 per cent of workers in the ORG data files have their earnings imputed using a "cell hot deck" method. The authors argue that this can create downward biases when the attribute being studied (e.g. union status) is not a criterion used in the imputation itself. In addition, there will be biases in other variables, such as schooling, where we understand only a limited number of controls (4) get used for the imputation even though sixteen education categories are available.

Furthermore, the proportions of the wage data that are imputed have increased in importance since a sample redesign of the CPS in 1993.[6] Unfortunately, it is not a simple matter to exclude -- in a way consistent over time -- the individuals who have imputed earnings. If past values of wages and/or other variables in a state are used to impute current wage values (as we understand is the case), this is likely to bias upwards the coefficient on the lagged dependent variable whenever the imputed observations are included. In other words, this would tend to distort research towards finding a Phillips curve.

We attempt to make an adjustment for this problem. We exclude the 'imputed' earners in all years except for 1994 and 1995. In those years it is not possible to make sample exclusions because allocation flags are unavailable in 1994 and only for a third of the months in 1995 (September-December). Allocation flags appear to be unreliable for 1989-93, moreover, when the proportion of earnings that were allocated mysteriously drops from 14.4% in 1988 to approximately 4% for the five years. A significant number of individuals in 1989-1993 had

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[6] The numbers of wage observations followed by the percentage imputed in parentheses (hourly + non-hourly paid) in the NBER MORG are given below based on the variable *I25d*. In 1995 the allocation information is only available on one-third of wage observations -- hence the small sample. In 1994 it is not possible to identify any individuals with imputed earnings.

1979 171,745 (16.5%)	1988 173,118 (14.4%)	1997 154,955 (22.2%)
1980 199,469 (15.8%)	1989 176,411 (3.7%)	1998 156,990 (23.6%)
1981 186,923 (15.2%)	1990 185,030 (3.9%)	1999 159,362 (27.6%)
1982 175,797 (13.7%)	1991 179,560 (4.4%)	2000 161,126 (29.8%)
1983 173,932 (13.8%)	1992 176,848 (4.2%)	2001 171,533 (30.9%)
1984 177,248 (14.7%)	1993 174,595 (4.6%)	2002 184,137 (30.4%)
1985 180,232 (14.3%)	1994 170,865 (0%)	2003 180,830 (32.0%)
1986 179,147 (10.7%)	1995 55,967 (23.3%)	
1987 180,434 (13.5%)	1996 152,190 (22.2%)	

earnings allocated but the Bureau of the Census apparently is unable to provide further details. It is a concern that these imputations of earnings do not appear to be random.

Our wage equations typically allow for a lagged dependent variable, w_{t-1}, alongside a set of standard variables that include age, gender, schooling, and industry variables. The equations also include year dummies. Other independent variables are the log of the local unemployment rate, union density, and average weekly benefits. These are all available at the state level. State union density rates by year were obtained from Union Membership and Earnings: Data Book, published by the Bureau of National Affairs, Inc., written by Barry Hirsch and David Macpherson (2002), and are downloadable at http://www.unionstats.com/. Average weekly UI benefits data were obtained from the Unemployment Insurance Financial Data Handbook published by the U.S. Department of Labor Employment & Training Administration. Data are downloadable at http://atlas.doleta.gov/XDMS/indexfrm.xml. The most recently available information on UI benefits is for 2001.

Most states have benefit levels that are intended to replace a little under one half of lost wages. However, there are maxima and minima on payments to unemployed individuals, and these are the source of much of the state-by-state variation in the data.

The maxima guarantee that many high-wage workers will receive less than half their average lost earnings; the minima mean that some low-wage workers earn more than fifty percent of their lost wages. There are considerable differences by state in the benefits maxima. In 1995, for instance, these ranged from a low of $175 per week in Missouri to a high of $362 in New Jersey (O'Leary and Rubin, 1997, page 174). The average wage replacement rate since 1938 has been approximately 35% (O'Leary and Rubin, 1997).

As an illustration, the Appendix gives data on benefit levels and replacement rates for the year 2000. The highest average level of benefits was in Massachusetts ($293) and the lowest in Puerto Rico ($104) and Mississippi ($157). Replacement rates -- defined as the proportion of workers' income replaced by benefits -- were greatest in Hawaii (50%) and lowest in California (20%).

5. Results on Hourly and Annual Wages and Earnings

In the US there was an early literature which found evidence of a positive relationship between wages and the unemployment rate. This includes papers by Hall (1970, 1972), Reza (1978), Behman (1978), Roback (1982), Adams (1985), Marston (1985) and Topel (1986). In Blanchflower and Oswald (1994a) we showed that such results were sensitive to the inclusion of region fixed-effects and to the time period used.

If few or no controls are used, the state-level correlation in the United States between wage levels and unemployment rates is positive. This can be viewed as a version of Harris-Todaro compensating differentials. As an illustration, on the latest US data, and after pooling the years 1979-2002, we obtain the following simple forms for log hourly earnings equations (where state unemployment is U):

	(1)	(2)	(3)
Log U	+.0333	-.0188	-.0373
	(2.71)	(2.72)	(1.92)
LR average Log U			+.1181
			(4.67)
State dummies (50)	No	Yes	No
Adjusted R^2	.8529	.9819	.8554
N	1223	1223	1223

These three equations include a set of year dummies (1979-2002); t-statistics are in parentheses.

Here it can be seen that the cross-section correlation between pay and joblessness is indeed positive, with an elasticity of 0.0333, when state dummies are omitted (there are no controls in this equation: the estimation is simply on year-state cell averages). Yet, in the spirit of a wage curve, the introduction of state dummies, in the second column, overturns the Harris-Todaro positive slope. The gradient changes to -0.0188 with a t-statistic of 2.72. A similar finding is produced in the third column where, instead of state dummies, the long-run average unemployment rate over the period (denoted LR average Log U) is entered as a regressor. There is then evidence of both a negative effect from the level of current unemployment and a positive effect from the long run average of unemployment. These elementary equations are not, of course, complete specifications. They are designed instead to illustrate how a Harris-Todaro positive correlation between area unemployment and area wages comes to be reversed.

Table 1 reports wage equations for the states of the US between 1980 and 2001. Our data set provides a sample size of 1122 state-year observations.[7] Column 1 of Table 1 is provided as a benchmark or extreme. Year dummies are included. State dummies are omitted. This column 1 specification gives the impression that an approximate Phillips Curve operates: the coefficient on lagged wages is 0.955. In other words, the equation is close to being one that could be rewritten with a pure change in log wage as the dependent variable (which would be literally correct if the coefficient on the lagged dependent variable were 1.0). The control variables in column 1 are age, age squared, and the percentage male -- all measured as state averages -- along with twenty year-dummies to pick up U.S.-wide changes in the labor market. A union density variable is also included, which enters positively and significantly.

It would be a mistake, however, to conclude that there is a simple regional Phillips Curve in these data. The addition of a set of state fixed-effects and industry controls has a marked effect (the industry controls are the proportion of workers in a 2-digit industry in a state-year cell). As can be seen, in column 2 of Table 1, the coefficient on the lagged dependent variable drops to 0.6990. This means, solving out for the steady-state level of wages, that the long run unemployment elasticity of wages is now estimated at -0.08. This number is obtained by dividing –0.0246 by 1 – 0.6990.

[7] Year 1979 has to be omitted because it is needed to generate the lagged dependent variable.

This is evidence of a U.S. wage curve. At -0.08, its elasticity is the same as in the wage-curve literature summarized above. Nevertheless, Column 2 of Table 1 exhibits a considerable amount of autoregression in pay, and Blanchard and Katz (1997) and others make a valuable point in drawing attention to some Phillips-like effects in regional wages in the United States.

Although our paper goes on to consider many variations and checks, there is a sense in which the main point emerges from this simple comparison of the first two columns of Table 1. Column 1 appears to be a Phillips Curve, but is specified too simply. Column 2, a fuller specification, is consistent with the kinds of wage curves that have been found in other countries. However, compared with some nations, the coefficient on the lagged dependent variable in the U.S. is high.

The level of state unemployment benefit enters positively and significantly in these, and later, wage equations. As far as we are aware, this finding is a new one. The same result holds true if the b variable is entered as a benefit/pay replacement ratio (this ratio can be obtained from the same data source and is defined as the average weekly benefits/average weekly wages in the state from UI records). Ideally there should be an instrument for b, but here we have to treat benefits as exogenous. Table 2's second column also finds that higher union density in a state is associated with higher wages in that state, ceteris paribus.[8]

Column 3 of Table 1 switches to include the log, rather than the level, of benefits. The change makes little difference. The size of the effects from benefits in column 3 seems plausible: a 100 per cent increase in the amount of benefit increases hourly wages in the long run by approximately 15 per cent.

In line with the theoretical framework outlined earlier, it is necessary to address the endogeneity of wages and unemployment. Column 4 of Table 1 begins this. It is estimated by Two Stage Least Squares (2SLS) and instruments the log of the unemployment rate with one and two lags of itself. Instrumenting produces a noticeable change in the size of the effects. Comparing Column 4 to Column 2, the coefficient on the lagged dependent variable is approximately the same, but the coefficient on unemployment has almost doubled in absolute value. As a result, the estimated value of the long run unemployment elasticity of pay rises in absolute value to close to -0.16.

Adding in state-specific time trends, as done by Bell, Nickell and Quintini (2002) for Great Britain, turns out to reduce the coefficient on the lagged dependent variable still further. This is done in Column 5 of Table 1. Here, although still statistically well determined, the estimated size of the unemployment elasticity of pay is only approximately -0.03. However, no instrumenting is done in column 5.

Within Table 1, Columns 6 and 7 provide the fullest specifications. The unemployment variable is instrumented; an unemployment benefits measure is included as an independent variable; union density is included; there are year dummies, industry controls, state dummies, and state time trends. In both columns, the unemployment elasticity of pay is estimated at approximately -0.1. Given their completeness, it might be natural to view these as the most compelling specifications.

[8] See also, for example, Hirsch and Schumacher (2004) and Blanchflower and Bryson (2003, 2004).

An alternative to estimation with year dummies is to include instead the aggregate wage in the United States (see Blanchard and Katz, 1999). This turns out here to make relatively little difference. Estimates of this kind using the private sector hourly wage of production workers, based around a version of Table 1, are given in Appendix 2a using both OLS (column 1) and GLS (column 2). The results are the same. We also experimented with the private sector weekly wage of production workers and obtained similar results. For completeness, to allow for the possibility of endogeneity of the benefits variable, b, columns 3 and 4 in Appendix Table 2a report a version of Table 1 with that variable omitted; there is still evidence of a wage curve whether OLS or GLS is used (columns 3 and 4 respectively).

How robust is the wage curve?

We now attempt to probe the robustness of the -0.1 finding.

Tables 2 and 3 perform simple experiments. In each table the sample is first restricted to Right-to-Work (RTW) and then to non-RTW states.[9] In effect, this is a division between states where unionism is weak and where it is stronger. Second, the sample is split into halves, and equation estimates are reported separately for the periods 1980-1991 and 1992-2002. Third, education controls are added. Table 2 uses OLS; Table 3 instruments the unemployment rate with its first two lags.

Interestingly, the coefficient on the lagged dependent variable in Table 2 is lower in RTW states than it is in non-RTW states. This may in part be because union contracts are traditionally for a period of two or three years, so in this sector there is automatically a degree of autoregression. In Table 2 the long run unemployment elasticity of wages changes a little: it is –0.05 in Column 1 and –0.08 in Column 2. Barth et al (2002a, b) argue that wage flexibility is greater in the non-union sector in the US. Indeed, they state that there is no wage curve for their sample of union workers. However, their work had to rely on fairly small cell-sizes in the union sector. This is likely to be a particular problem in the southern RTW states where union membership is low.

Adding education controls -- again these are averages at the state level -- into Table 2 reduces the unemployment elasticity of pay to a value of –0.04 in Column 3. There is, however, no completely consistent way to control for educational levels; there was a change in the CPS education question from 1992.

When the data are divided into two time periods, an interesting outcome is produced. This division is done in Columns 3 to 6 in Table 2. The lagged dependent variable is found to have a markedly smaller coefficient in the second period (that is, for the years 1992-2001) than in the first (period 1980-91). Moreover, in Column 4, the coefficient on the local unemployment rate is almost exactly zero. However, these estimates ignore the potential problem of simultaneity, and the statistical significance of the joblessness variable returns when the variable is instrumented.

[9] The Right-To-Work (RTW) states are Alabama, Arizona, Arkansas, Florida, Georgia, Idaho, Iowa, Kansas, Louisiana, Mississippi, Nebraska, Nevada, North Carolina, North Dakota, Oklahoma, South Carolina, South Dakota, Tennessee, Texas, Virginia, Wyoming and Utah. For details see http://www.nrtw.org/rtws.htm.

Endogeneity of unemployment is allowed for in the columns of Table 3. Across specifications, the long-run unemployment elasticity of the wage curve varies from a low of –0.06 to a high of –0.14. The fullest specification, and probably the most reliable, is in Columns 5 and 6. State time trends are included. So too are year dummies, industry controls, state fixed-effect dummies, and in Column 6 of Table 3 also education controls. In the first period, the unemployment elasticity of pay is estimated at -0.07. In the later period, the elasticity is -0.12. It seems best simply to conclude that these estimates cluster around –0.1.

Table 4 now explores the sensitivity of the results to measurement error by instrumenting the lagged dependent variable itself. The approach follows that of Staiger, Stock and Watson (2002). To obtain an instrument, the authors' paper exploits the fact that the ORG files are obtained monthly.

A suitable instrument must be correlated with lagged pay but uncorrelated with the other variables in the wage equation. Our approach relies on the fact that the data set is collected throughout each calendar year. We create two wage variables. One averages the pay data in even months (i.e., February, April, June and so on). The other is an average of pay in the odd months. This is done to overcome the problem that, as many states have small numbers of CPS respondents, estimated wage equations can exhibit errors-in-variables bias.

Staiger et al (2002) split the monthly ORG files into these two independent samples (although households appear twice in the survey, the odd and even months have no households in common). Their innovative suggestion is this:

> "Estimates from both the odd and even month samples will be measured with error, but because the samples are randomly drawn, the estimation error is independent in these two samples. Thus one set of estimates can be used as an instrument for the other set of samples." (p.39, 2003).

A number of possibilities exist, so we go further and obtain separate wage estimates for individuals in the CPS in (a) the odd months, (b) the even months, (c) the first six months, (d) the second six months, (e) the first and third quarters, and (f) the second and fourth quarters. All should be valid instruments. A further question that needs to be addressed is whether it is appropriate to instrument the odd months with the even, as Staiger et al (2002) do, or vice versa.

Table 4 presents various specifications. The most persuasive equations are in part C of the table, but the others are presented for completeness. Table 4 has three parts: the first, A, where industry dummies are not included as controls; part B where they are; and part C which includes state time trends and industry controls. Interestingly, the lagged dependent variable has a bigger coefficient when the evens are used to instrument the odds than when the reverse is true. A similar picture emerges when the two halves or the quarters are used; the use of past values to instrument current values generates a smaller coefficient on the lag on the dependent variable.

Table 4 deliberately presents a battery of estimates to show how the results move away from a lagged dependent variable approximately equal to unity. If the past affects the future only through the present, then instrumenting the present with the past would seem appropriate and the cleanest way to do so is to instrument the second six months with the first six months. Our preferred specification in Table 4 is therefore row 4 in part C where the even months are instrumented with the odd months. Here the lagged dependent variable has a coefficient of 0.60 and the estimated long run unemployment elasticity of wages is approximately -0.075.

Table 5 goes a little further. It reports results where the lagged dependent variable w_{t-1} and the unemployment rate U_t are both instrumented. Table 5's aim is to try to correct, within the same estimating equation, for the endogeneity of unemployment and measurement error in the dependent variable. In the equations of Table 5, the dependent variable is derived from the second six months of CPS data, January-June. Then the lagged dependent variable is itself instrumented with the wage derived by averaging data from the first six months, July-December. As before, the unemployment rate in time t is instrumented with the unemployment rates at t-1 and t-2.

The coefficient on the lagged dependent variable is approximately 0.81 in the first column of Table 5. It has a small standard error, and a (Phillips Curve) null hypothesis of unity is still comfortably rejected at normal significance levels. Column 2 of Table 5 adds state time-trends. Once this is done, the coefficient on w_{t-1} drops to approximately 0.71. As before, there appears to be some difference between the right-to-work (RTW) states and the non-RTW states. The coefficients on the lagged dependent variable are noticeably different between columns 3 and 4 and 5 and 6. The addition of state time-trends has a noticeably large impact in the sample of RTW states, whereas in non-RTW states the results in columns 5 and 6 are similar to one another. The long run unemployment elasticity of wages is somewhat lower in the RTW states.

Another experiment is to alter the nature of the dependent variable. Up to this stage, the estimation in the paper has used data on hourly wages. It is a matter of judgment whether a pay variable is better defined as hourly or annual. While theoretical economics might favor an hourly variable, there are practical disadvantages in empirical work. Most American workers are not paid by the hour and are not aware of their own hourly wage rate but they do know their annual pay. Annual earnings data are reported in the CPS in March of each year. Annual earnings are reported on a worker's W2 form from their employer, which also appears on the individual's tax return filed in April of each year.

From Blanchflower and Oswald (1994a), Blanchard and Katz (1997), and Card (1995), it is known that wage curves estimated with annual pay variables tend to exhibit relatively little autoregression. To explore this in the latest data, Table 6 switches to annual earnings as a dependent variable. The table draws upon data from the Annual Demographic March files of the CPS. These files contain information on each interviewed worker's earnings in the preceding year. We calculate weighted averages of earnings, using the *earnwt* variable, and map that to our file. This is done at the level of the state/year cell. Other control variables, including age, gender and industry, are derived from the ORG files and are the same as those used previously in Tables 2-4. Data are available on the numbers of weeks worked during the preceding year and the usual hours. In Table 6, columns 1-4, 6 and 7 are pay equations estimated for full-time/full-year workers, defined as those who worked for 52 weeks and at least 35 hours per week. In columns 5 and 8 of Table 6 we impose no such restrictions and use the level of wages in the region. Sample sizes for state/year cells range from an average of 760 workers in Vermont to 6,527 in

California. Overall, sample sizes are smaller in the March files than in the ORGs; they average about 75,000 cases in the former case and 165,000 in the latter. To deal with the possibility that small sample sizes might be responsible for the results, in columns 4 and 5 a set of weighted estimates are also given. These use as weights the cell counts obtained when we averaged. Reassuringly, weighting turns out to have a negligible effect.

Table 6's results are fairly strong. In all but one column there is a well-determined effect from local unemployment upon the level of pay. The estimates of the long-run unemployment elasticity in Table 6 are generally consistent with, though a little lower than, those obtained in Tables 2-5. They range here from -0.02 to -0.12. As in previous research, the coefficient on the lagged dependent variable is smaller than in the ORGs. Including state effects, year effects and industry controls, the coefficient on w_{t-1} is around 0.35. Adding in state time-trends reduces the lagged coefficient on the wage still further, to around 0.13, whether OLS or GLS is used. Average weekly benefits remain strongly significant at conventional levels, and positive in all specifications.

Appendix Table 2b explores a number of other issues.

First, it examines the extent of any bias when estimation does not take account of the fact that a significant proportion of wage observations in the ORGs are imputed. Column 1 of this Table is the equivalent of column 2 of Table 1; column 2 of Appendix Table 2b takes the same specification but now includes all observations of the wage, whether imputed or not. The bias from including the imputed observations appears to be miniscule.

Second, we use an alternative estimation method. Developed by Blanchard and Katz (1997), it involves a two-stage procedure. At the first stage a wage equation is estimated in each year with a number of controls included to distinguish differences in worker and workplace characteristics, and the coefficients on the state dummy variables are then used at the second stage. In five of the six specifications that Katz and Blanchard (1997) report in their Table 2, they do not include any average worker or workplace controls at the second stage; when they do, the size of the lagged dependent variable falls from .633 to .258 when the March CPS (1980-1991) is used. Column 3 of Appendix Table 2b provides the results using their two-stage procedure with the same set of controls as used in column 2 also with the imputed data excluded for the period 1980-2001. Now the lag is .744 with industry controls included, and comparable to the result in column 1, with a long-run elasticity of -.11. In column 4, the coefficient on the lagged wage is higher when the industry variables are excluded. Omitting the age and gender variables, in column 5 of Appendix Table 2b, has little further impact.

If the sample is restricted to the years 1980-1991 that were available to Blanchard and Katz, and we include only non-imputed data at the first stage, and only year and state dummies as additional controls at the second stage, we get exactly the result they obtained in column 3 of their table. As can be seen below, the results are different when either more years or more controls are added at the second stage or both (controls are as in Appendix Table 2b, column 2).

	Lagged wage	Log U	Long run elasticity
Blanchard Katz 1980-1991 no controls	.91	-.04	-0.444

1980-1991 no controls	.91	-.04	-0.444
1980-1991 with controls	.77	-.02	-0.087
1980-2001 no controls	.88	-.03	-0.250
1980-2001 with controls	.74	-.03	-0.115

Third, Doug Staiger has suggested to us that if wages genuinely follow a random walk, and there are no state-level effects on wages from variables like climate or regional prices, then any specification with state-specific parameters (e.g. state fixed effects or state time trends) will be biased toward finding a wage curve. An analogy is to running spurious regressions in time series analysis. Supposing that Y follows a random walk:

$Y_t = Y_{t-1} + e_t$,

then, if we run the regression,

$Y_t = b_0 + b_1 Y_{t-1} + e_t$

the coefficient on b_1 is biased downward below 1, and the usual OLS test of $b_1=1$ over-rejects. This problem is worse if we include a time trend on the right hand side. In the state data, the analogy is that when we include state dummies or state-specific time trends, the coefficient on the lag is biased down. The size of such bias is uncertain. Columns 6 and 7 of Appendix Table 2b are two simple ways of addressing this, but are suggestive only. Column 6 replaces the state fixed-effects with the long-run average of the log unemployment rate for the period 1979-2003, while column 7 uses a 19 year moving average of state unemployment rates -- a 19 year average is used because we have data back only as far as 1960. These variables enter positively, as we found earlier, suggesting a version of Harris-Todaro compensating differentials. The lag on the dependent variable is now .839 with a long-run elasticity of -.17 and -.15 respectively. Evidence of a US wage curve remains.

Unemployment equations

Lying behind each of these wage equations is, implicitly, an unemployment equation. As explained earlier in the paper, it takes the form $U_t = \Omega(U_{t-1}, U_{t-2}, w_t, \text{controls})$.

The natural justification for this form is to define area unemployment as $U = P(w_t, \text{controls}) – N(w_t, \text{controls})$ where P is the participation rate in the area and N is the level of employment demand in the area. We assume the customary AR(2) form of an employment function, and also that participation is a non-decreasing function of the wage in the area. Then the unemployment rate in a state, in time period t, is a function of itself lagged twice and of the current wage rate. With these assumptions, the wage should enter positively in the estimated unemployment equation. The greater is the price of labor, the lower will be the level of labor demand and the higher the degree of labor force participation, and so, other factors constant, the higher will be the unemployment rate.

In principle the wage variable inside the unemployment equations should be in real rather than nominal terms, which means that annual state-level price indexes are required. Such data do not exist. However, subject to the assumption that there are not significant variations in the state structure of prices, year-dummies can be used as an alternative.

It is useful to study the exact empirical structure of this unemployment equation when fitted to US data over the period. Table 7 does so. The table models the log of the unemployment rate in the state/year cell for the period 1980 to 2002. This effectively treats US states as a panel.

A central concern of the analysis here is whether well-behaved unemployment equations can be estimated once both state fixed-effects and year effects are incorporated. There appears to have been little attempt before in the U.S. literature to estimate such equations. The well-known paper by Blanchard and Katz (1992) on regional evolutions comes close to this, but it focuses primarily on employment rather than unemployment data per se.

Table 7 presents a selection of state unemployment equations. It includes controls for workers' age and its square, gender, the log of hourly and annual pay, and union density in the state.

Extra controls are added as we move across the columns of Table 7. In later columns, in line with the theoretical framework, the level of pay is instrumented using the state benefits level. Column 1 contains only year dummies. The coefficient on the first of the two lagged dependent variables is approximately 0.9. Hourly pay is positive but poorly determined. Union density has no effect. Column 2 of Table 7 adds state fixed-effects into the specification. Once this is done, the unemployment equation begins to take the form that might be expected from economic theory. Unemployment is autoregressive: U_{t-1} enters with a coefficient of approximately 0.8 and U_{t-2} with a coefficient of approximately -0.08. Log hourly pay now has a large and well-determined positive coefficient. The coefficient on union density is positive, although the t-statistic remains weak. These results are consistent with an inverted labor demand function.

It is natural to examine the implied long run elasticity. Column 2 of Table 7 can be solved out, in the usual way, for a steady-state logarithmic equation in which the wage elasticity of unemployment is approximately 0.6.

Column 3 of Table 7 introduces 44 industry dummies. The size of the union density effect increases and becomes significantly different from zero at the 5% level. The second lag on unemployment loses significance. As an experiment, column 4 replaces the hourly pay variable from the ORGs with the annual pay variable taken from the March CPS data. The pay variable remains significant at the 5% level.

In Table 7, columns 5 through 7, the unemployment equations are estimated with 2SLS. The pay variables are instrumented by the state's average weekly benefit. This correction for endogeneity has the effect of markedly increasing the coefficient on pay in an unemployment equation. It also reduces the size of the U_{t-2} effect. The coefficient on union density is then poorly determined in all later specifications. Layard, Nickell and Jackman (1991) and OECD (1994) have argued that high levels of unionization are part of the explanation for differences in unemployment across countries.[10] Whatever the evidence across countries, it does not seem, controlling for the wage,

[10] For more on the causes of unemployment across OECD countries, see Blanchflower (2001) and Oswald (1997b).

that there is reliable evidence for this claim across the states of the USA. However, we have not attempted to explore the reduced-form effects of unions.

6. Empirical Results on Happiness and Job-Finding

This section presents evidence that can be viewed as complementary to the earlier wage equations. It uses subjective data to explore how workers' wellbeing and perceptions are influenced by the local rate of unemployment.

Implicit in the paper's analytical approach to the labor market are two building blocks. One is the idea that a rise in local unemployment weakens workers' utility relative to employers' utility. To be more specific, the expected utility level of a worker should, according to the theory, be lower in areas where unemployment is high. The reason is that in such areas the typical employee is more fearful of job loss. Such a prediction is captured in the theoretical equation 1 given earlier in the paper (because the level of expected utility is a decreasing function of unemployment, U). Second, the worker's perceived chance of finding another job if laid off is, according to the theory, a declining function of the local unemployment rate. In principle, both these ideas can be checked empirically. The problem in practice is the availability of appropriate data.

We test these two hypotheses by using information on reported happiness and on workers' job-finding perceptions. We draw upon data from the General Social Survey (GSS) series for the United States. Although still little used by economists, the GSS contains useful data on workers' attitudes and subjective wellbeing.

Happiness data in the GSS are available for most of the years between 1972 and 2002 (there were no surveys conducted in 1979, 1981, 1992, 1995, 1997, 1999 and 2001). We were able to map onto the survey series the state/year unemployment rates.[11] Data are available in 45 states; surveys have not been conducted in 6 small states (Maine, Hawaii, New Mexico, Nebraska, Idaho and Nevada).

First, we estimate a series of happiness equations based upon the specifications used -- for a different purpose -- in Blanchflower and Oswald (2004). The specific focus here is upon workers' subjective wellbeing and on how that is shaped by the state of the local labor market. Second, we explore a variable measuring an individual's fear of unemployment, and again attempt to see how it is affected by labor-market conditions.

In the GSS, respondents are asked to report their levels of subjective wellbeing. The question asked is:

> *Taken all together, how would you say things are these days -- would you say that you are very happy, pretty happy, or not too happy?*

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[11] We are grateful to Tom W. Smith at NORC for giving us permission to use, and providing us with, the state level codes that allowed us to map in the state/year unemployment rates.

On average, between 1972 and 2002, approximately 32% of respondents said they were very happy, 56% were pretty happy, and 12% not too happy. We treat this as a dependent variable.

The first two columns of Table 8 estimate ordered logit happiness equations for the pooled years 1972-1998. State dummies and year dummies are included. Because of the inclusion of the state unemployment rate in an equation estimated at the level of the individual, standard errors are adjusted for clustering at the state level.

For the sake of completeness, the first column of Table 8 estimates a happiness equation for the full GSS sample, that is, one including non-workers. The log of the local unemployment rate here enters negatively, with a t-statistic of 1.78. In other words, Americans who live in a state with higher unemployment report lower happiness scores, but this effect is not quite statistically significant at the 5% level.

When the sample is restricted to workers, however, in column 2 of Table 8, it can be seen that the unemployment effect becomes better determined. The coefficient is now –0.17 with a t-statistic of 2.19. Although the coefficients in ordered logits are not straightforward to interpret, it can be checked that this implies that higher unemployment in the area makes employees systematically less likely to report themselves in the two happier categories of the GSS survey responses. This is consistent with the conceptual framework sketched earlier.

In Table 8, therefore, where pay is deliberately omitted from the equation, a higher local unemployment rate is associated with a lower reported level of workers' 'utility'. This finding is similar to that in country-level data, as discussed in Di Tella, MacCulloch and Oswald (2001) and Oswald (1997a). It also seems consistent with UK results, of a different kind, on 'fear' of unemployment in Blanchflower (1991).

Column 3 of Table 8 estimates an ordered logit where the dependent variable is employee responses to the following question.

> *"About how easy would it be for you to find a job with another employer with approximately the same income and fringes you now have? Would you say very easy, somewhat easy or not easy at all?"*

Responses vary cyclically over time and, consistent with common sense, appear to be negatively correlated with the aggregate unemployment rate.

	Not easy	Somewhat easy	Very easy	Unemployment rate (%)
1977	43	30	27	7.1
1978	39	33	28	6.1
1982	53	26	21	9.7
1983	51	29	19	9.6
1985	43	32	25	7.2
1986	39	33	28	7.0
1988	35	37	28	5.5
1989	38	28	34	5.3

1990	38	30	32	5.6
1991	40	36	24	6.8
1993	45	33	22	6.9
1994	46	33	21	6.1
1996	40	33	27	5.4
1998	33	36	31	4.5
2000	29	33	38	4.0
2002	36	37	27	5.8

Data on this survey question are available only for the years 1977, 1978, 1982, 1983, 1985, 1986, 1989-1991, 1993, 1994, 1996, 1998, 2000 and 2002. The percentage of people saying it would be 'very easy' to find a job was highest in 2000, when the US unemployment rate was lowest at 4%. Conversely, the proportion saying it was 'not easy' was highest in 1982, when unemployment was at its peak of 9.7%.

Column 3 of Table 8 estimates an ordered logit where the dependent variable is the individual's perception of how easy it would be to find a comparable job elsewhere. This is a subjective variable, of course, but it seems potentially of interest. The estimated perceived-chance-of-finding-a-job equation is an attempt to show how fear of unemployment (or more precisely, the local rate of unemployment) works: it influences workers' feelings about the probability of finding a new job. Such a finding is required if the conceptual framework proposed for the wage curve is the correct one. Standard errors are once again adjusted for clustering. The likelihood-of-finding-a-job equation captures the subjective probability of getting work and has a negative gradient with respect to the state unemployment rate. With a t-statistic of 4.79, the impact of the log of state unemployment in column 3 of Table 8 is a well-determined one. This finding is consistent with one part of the wage-curve framework.

The fact that the log of state unemployment enters significantly and negatively is compatible with results in Blanchflower and Oswald (1999) and Schmidt (1999). A greater rate of joblessness in the worker's region leads, it might be said, to a diminution of that worker's power relative to the firm. As local unemployment rises, workers become less confident about being able to secure a position with another employer.

Although the data used in this section are of a more subjective kind than used by most labor economists, they paint a picture that is consistent with the analytical foundations of the wage-curve. Unemployment in the local area acts to lower workers' reported happiness and to reduce an individual's perceived chance of finding a job.

7. Conclusion

Economics has few empirical laws. It is important to know if the wage curve qualifies as one. Although evidence for such a curve has been found in more than 40 countries, the existence of a wage curve in United States data has been viewed as more controversial.

The estimates in this paper use a longer period of data than has been possible before. It also attempts to correct for endogeneity and measurement error. The paper finds -- consistent with an efficiency-wage framework -- that wages are lower in U.S. states with higher unemployment. As has been discovered in many other nations, this wage curve has a long run elasticity of approximately –0.1.

A number of other results emerge. They are compatible with the theoretical framework proposed in the paper.

- Wages are greater in states with more generous unemployment benefits.

- The higher is the unemployment rate in a state, the lower are the levels of happiness reported by workers in that state.

- The higher the unemployment rate in a state, the lower is the typical worker's perceived chance of finding a job.

- Evidence for the wage curve is particularly strong in non-union (Right-to-Work) states.

- Except in isolated specifications, there is not persuasive support for a simple Phillips curve. It seems more sensible to view the data as being characterized by dynamic fluctuations around a long-run wage curve. State wages in the United States are autoregressive (as explained in Card and Hyslop, 1997, and Blanchard and Katz, 1997), but not extremely so. In many specifications, the coefficient on a lagged dependent variable is 0.5 or smaller.

This paper's results seem consistent with the empirical law proposed in Blanchflower and Oswald (1994a). The United States has a wage curve.

Table 1. US State Log Hourly Wage Equations, 1980-2001

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Log hourly pay$_{t-1}$	.9553	.6990	.6788	.7143	.4921	.5390	.5331
	(106.88)	(33.61)	(31.98)	(33.56)	(19.09)	(19.51)	(19.12)
Log unemployment rate$_t$	-.0229	-.0246	-.0245	-.0445	-.0134	-.0418	-.0405
	(9.97)	(6.50)	(6.56)	(8.01)	(3.37)	(6.55)	(6.37)
Average weekly benefit$_t$	.00004	.0002		.0002	.0004	.0004	
	(1.30)	(4.03)		(4.69)	(5.46)	(5.90)	
Log weekly benefit$_t$			.0495				.0642
			(5.70)				(6.12)
Union density$_t$	.0004	.0013	.0011	.0015	.0015	.0016	.0016
	(3.38)	(2.84)	(2.80)	(3.36)	(3.39)	(3.50)	(3.38)
Constant	-.0303	.0719	-.1498	.3718	.1591	.6137	.0049
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry controls	No	Yes	Yes	Yes	Yes	Yes	Yes
State fixed effects	No	Yes	Yes	Yes	Yes	Yes	Yes
2SGLS	No	No	No	Yes	No	Yes	Yes
State time trends	No	No	No	No	Yes	Yes	Yes
N	1122	1122	1122	1122	1122	1122	1122
Within R^2				.9957		.9963	.9964
Between R^2				.9664		.9055	.9031
Total R^2	.9953	.9911	.9965	.9906	.9965	.9762	.9765
F-statistic	8308	2278.63	3301.1		2223.4		
Wald/Chi^2				1.79e+07		2.00e+07	2.01e+07
Long run unem. elasticity	-.5123	-.0817	-.0763	-.1558	-.0264	-.0907	-.0867

Notes: the log of the unemployment rate is instrumented with its 1 year and 2 year lags. All equations also include the averages of age, age squared and gender in each state/year cell. Source: weekly wage benefit amount available from *Unemployment Insurance Financial Data Handbook* published by the U.S. Department of Labor Employment & Training Administration. Data are downloadable at http://atlas.doleta.gov/XDMS/indexfrm.xml for Handbook 394. State level union density rates are obtained from Union membership and earnings: data book by Barry Hirsch and David Macpherson and downloadable at http://www.unionstats.com/. All other data are weighted averages in each state/year cell obtained from the ORG files of the CPS 1980-2001. t-statistics are in parentheses.

Table 2. OLS US State Log Hourly Wage Equations, 1980-2001: Disaggregated by RTW States, Education and Periods

	(1)	(2)	(3)	(4)	(5)	(6)
Log hourly pay$_{t-1}$	.5548	.7014	.5988	.3555	.1992	.0259
	(14.35)	(24.59)	(18.36)	(8.75)	(4.79)	(0.52)
Log unemployment rate$_t$	-.0223	-.0240	-.0169	.0010	-.0161	.0110
	(3.26)	(4.68)	(3.06)	(0.13)	(3.09)	(1.29)
Average weekly benefit$_t$	.0002	.0002	.0003	.0003	.0006	.0003
	(3.07)	(3.36)	(3.93)	(3.02)	(4.99)	(2.01)
Union density$_t$	.0023	.0011	.0019	.0015	.0016	.0014
	(3.07)	(1.92)	(3.69)	(1.60)	(3.11)	(1.46)
Constant	-.6479	.2929	-.3785	.9348	-.1369	.6289
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry controls	Yes	Yes	Yes	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
State time trends	Yes	Yes	No	No	Yes	Yes
RTW state	Yes	No	N/a	N/a	N/a	N/a
Years of education	No	No	Yes	No	Yes	No
Education controls	No	No	No	Yes	No	Yes
1980-1991	No	No	Yes	No	Yes	No
1992-2001	No	No	No	Yes	No	Yes
N	484	638	612	510	612	510
Total R^2	.9967	.9966	.9943	.9923	.9962	.9947
Long run unemployment elasticity	-.0500	-.0804	-.0421	.0016	-.0201	n/a

Notes: see Table 1. The Right-To-Work (RTW) states are Alabama, Arizona, Arkansas, Florida, Georgia, Idaho, Iowa, Kansas, Louisiana, Mississippi, Nebraska, Nevada, North Carolina, North Dakota, Oklahoma, South Carolina, South Dakota, Tennessee, Texas, Virginia, Wyoming and Utah. For details see http://www.nrtw.org/rtws.htm. t-statistics are in parentheses.

Table 3. 2SLS US State Log Hourly Wage Equations, 1980-2001: Disaggregated by RTW States and Periods

	(1)	(2)	(3)	(4)	(5)	(6)
Log hourly pay$_{t-1}$	.4057	.5791	.7407	.3800	.2722	.0999
	(8.08)	(15.44)	(22.92)	(8.86)	(5.88)	(1.33)
Log unemployment rate$_t$	-.0549	-.0395	-.0352	-.0351	-.0514	-.1129
	(3.60)	(4.54)	(3.47)	(2.17)	(5.49)	(1.65)
Average weekly benefit$_t$	.0004	.0003	.0003	.0003	.0006	.0003
	(3.25)	(3.77)	(3.32)	(2.74)	(5.35)	(1.82)
Union density$_t$	.0027	.0013	.0018	.0012	.0020	.0007
	(3.37)	(2.03)	(3.41)	(1.24)	(3.68)	(0.54)
Constant	.2025	.8539	-.6104	1.3940	-.0821	2.8519
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry controls	Yes	Yes	Yes	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
2SGLS	Yes	Yes	Yes	Yes	Yes	Yes
RTW state	Yes	No	N/a	N/a	N/a	N/a
State time trends	Yes	Yes	No	No	Yes	Yes
Years of education	No	No	Yes	No	Yes	No
Education controls	No	No	No	Yes	No	Yes
1980-1991	No	No	Yes	No	Yes	No
1992-2002	No	No	No	Yes	No	Yes
N	484	638	612	510	612	510
Within R^2	.9969	.9966	.9902	.9841	.9935	.9834
Between R^2	.7904	.8895	.9425	.8740	.7296	.0944
Total R^2	.9839	.9763	.9727	.9278	.8970	.1607
Wald/Chi^2	8.96E+06	1.16e+07	9.30e+06	1.15e+07	1.26e+07	9.57e+06
Long run unemployment elasticity	-.0924	-.0938	-.1358	-.0566	-.0706	-.1254

Notes: see Table 1. The Right-To-Work (RTW) states are Alabama, Arizona, Arkansas, Florida, Georgia, Idaho, Iowa, Kansas, Louisiana, Mississippi, Nebraska, Nevada, North Carolina, North Dakota, Oklahoma, South Carolina, South Dakota, Tennessee, Texas, Virginia, Wyoming and Utah. For details see http://www.nrtw.org/rtws.htm. Instruments for the log unemployment rate are one and two lags of itself. t-statistics are in parentheses.

Table 4. Instrumenting the Lagged Dependent Variable: A Selection of Unemployment Elasticities of Pay, 1980-2001.

	Lagged dep var	Log Unemployment	Weekly Benefits	Union	Long run elasticity
A) Without industry controls					
1. Even instrumented with odd	.9119	-.0294	.00004*	.0009*	-.3337
2. Odd instrumented with even	.9487	-.0414	.00005*	.0012	-.8070
3. First half instrumented with 2nd half	.9749	-.0336	.00002*	.0008*	-1.3386
4. 2nd half instrumented with 1st half	.8795	-.0380	.00008*	.0012	-.3154
5. Qrtrs 1 & 3 instrumented with qrtrs 2 & 4	.9460	-.0415	.00003*	.0015	-.7685
6. Qrtrs 2 & 4 instrumented with qrtrs 1 & 3	.9156	-.0292	.00006*	.0005	-.3460
B) With industry controls					
1. Even instrumented with odd	.8228	-.0209	.00006*	.0007*	-.1179
2. Odd instrumented with even	.8925	-.0361	.0006*	.0014	-.3358
3. First half instrumented with 2nd half	.9147	-.0235	.0003*	.0007*	-.2755
4. 2nd half instrumented with 1st half	.7896	-.0336	.0001*	.0014	-.1597
5. Qrtrs 1 & 3 instrumented with qrtrs 2 & 4	.8935	-.0360	.00006*	.0018	-.3380
6. Qrtrs 2 & 4 instrumented with qrtrs 1 & 3	.8374	-.0215	.00005*	.00035*	-.1322
C) With state time trends and industry controls					
1. Even instrumented with odd	.6512	-.0149	.0001*	.0012	-.0427
2. Odd instrumented with even	.7908	-.0336	.0001*	.0016	-.1606
3. First half instrumented with 2nd half	.8359	-.0202	.00001*	.0010*	-.1231
4. 2nd half instrumented with 1st half	.6081	-.0294	.00007*	.0016	-.0750
5. Qrtrs 1 & 3 instrumented with qrtrs 2 & 4	.7927	-.0309	.00003*	.0022	-.1491
6. Qrtrs 2 & 4 instrumented with qrtrs 1 & 3	.7120	-.0197	.00004	.0006*	-.0684

*=insignificantly different from zero

Notes: the mean wages were calculated using the *earnwt* variable and excluding any individual whose earnings were allocated using the *I25d* variable. The exclusion was not used in 1994 and 1995 as the wage allocation variable *I25d* is not available in 1994 and for only 1/3rd of the months in 1995. T-statistics are in parentheses.

Table 5. 2SLS US State Log Hourly Wage Equations, 1980-2001: Disaggregated by RTW States and Periods (Instrumenting Lagged Pay and Unemployment)

	(1)	(2)	(3)	(4)	(5)	(6)
Log hourly pay$_{t-1}$	.8069	.7071	.6253	.4999	.8274	.8223
	(21.55)	(11.44)	(8.01)	(4.25)	(15.51)	(8.01)
Log unemployment rate$_t$	-.0482	-.0573	-.0666	-.0749	-.0426	-.0488
	(6.20)	(6.42)	(3.49)	(3.22)	(4.28)	(3.71)
Average weekly benefit$_t$	.0001	.0001	.0002	.0002	.0001	.0001
	(1.66)	(0.93)	(1.18)	(1.04)	(1.00)	(0.84)
Union density$_t$	.0016	.0017	.0035	.0035	.0015	.0015
	(2.64)	(2.54)	(3.09)	(2.96)	(1.80)	(1.56)
Constant	.4766	.7657	.3117	.8058	.2973	-.3835
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry controls	Yes	Yes	Yes	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
RTW state	N/a	N/a	Yes	Yes	No	No
State time trends	No	Yes	No	Yes	No	Yes
N	1122	1122	484	484	638	638
Within R^2	.9910	.9919	.9917	.9924	.9922	.9929
Between R^2	.9727	.9447	.8105	.6601	.9733	.9200
Total R^2	.9868	.9815	.9781	.9676	.9877	.9808
Wald/Chi^2	8.87e+06	8.95e+06	3.74e+06	3.87e+06	5.53e+06	5.70e+06
Long run unemployment elasticity	-0.2496	-0.1956	-0.1777	-0.1498	-0.2468	-0.2746

Notes: see Table 1. The Right-To-Work (RTW) states are Alabama, Arizona, Arkansas, Florida, Georgia, Idaho, Iowa, Kansas, Louisiana, Mississippi, Nebraska, Nevada, North Carolina, North Dakota, Oklahoma, South Carolina, South Dakota, Tennessee, Texas, Virginia, Wyoming and Utah. For details see http://www.nrtw.org/rtws.htm. Instruments for the log unemployment rate are one and two lags of itself. The dependent variable is pay from the second six months instrumented with the first six months. t-statistics are in parentheses.

Table 6. US State Log Annual Pay Equations, 1980-2001

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Log annual pay$_{t-1}$	.8955	.5059	.3509	.3626	.3466	.3349	.1327	.1374
	(61.25)	(18.73)	(12.19)	(12.86)	(12.77)	(11.36)	(4.03)	(4.32)
Log unemployment rate$_t$	-.0131	-.0226	-.0150	-.0181	-.0505	-.0258	-.0266	-.0104
	(2.69)	(3.26)	(2.09)	(2.64)	(7.56)	(2.49)	(2.17)	(1.36)
Average weekly benefit$_t$	.0002	.0007	.0006	.0006	.0005	.0006	.0005	.0005
	(3.71)	(7.89)	(6.11)	(7.29)	(6.56)	(6.35)	(3.92)	(3.91)
Union density$_t$	.0005	.0011	-.0006	-.0005	.0008	-.0004	-.0004	-.0006
	(1.97)	(1.36)	(0.77)	(0.70)	(1.08)	(0.44)	(0.46)	(0.6
Constant	.9933	4.6231	5.6956	5.7507	5.5047	6.3837	8.8832	7.9293
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry controls	No	No	Yes	Yes	Yes	Yes	Yes	Yes
State fixed effects	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
2SGLS	No	No	No	No	No	Yes	Yes	No
State time trends	No	No	No	No	No	No	Yes	Yes
Full-time/full-year	Yes	Yes	Yes	Yes	No	Yes	Yes	No
Weighted by cell count	No	No	No	Yes	Yes	No	No	No
N	1122	1122	1122	1122	1122	1071	1071	1122
Within R^2						.9822	.9848	
Between R^2						.7301	.4623	
Total R^2	.9797	.9841	.9867	.9903	.9933	.9369	.9028	.9879
F-statistic/Wald/Chi^2	1952.0	2067.6	889.6	1254.3	1812.2`	9.13e+07	1.01e+08	
LR unempt. elasticity	-.1254	-.0457	-.0231	-.0284	-.0773	-.0388	-.0307	n/s

Notes: the log of the unemployment rate is instrumented with its 1 year and 2 year lags. All equations also include the averages of age, age squared and gender in each state/year cell. Source: weekly wage benefit amount available from *Unemployment Insurance Financial Data Handbook* published by the U.S. Department of Labor Employment & Training Administration. Data are downloadable at http://atlas.doleta.gov/XDMS/indexfrm.xml for years up to 1997 and subsequently at http://atlas.doleta.gov/XDMS/indexfrm.xml. State level union density rates are obtained from Union membership and earnings: data book by Barry Hirsch and David Macpherson and downloadable at http://www.unionstats.com/. All other data are weighted averages in each state/year cell obtained from the March CPS files 1981-2002. Data relate to the preceding year so data for 2001 are obtained from the March 2002 Annual Demographic file. Full-time full year includes only individuals who worked 52 weeks and usually worked >=35 hours per week. t-statistics are in parentheses.

Table 7. State Unemployment Equations, USA: 1980-2002.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Log unemployment $rate_{t-1}$	.9165	.8127	.7284	.7876	.8087	.7320	.7769
	(31.26)	(27.59)	(24.37)	(24.20)	(25.03)	(20.76)	(18.48)
Log unemployment $rate_{t-2}$	-.0394	-.0758	-.0444	-.0999	.0119	.0545	-.0209
	(1.35)	(2.51)	(1.45)	(3.06)	(0.33)	(1.31)	(0.47)
Log hourly pay_t	.0445	.4268	.5470		1.5441	2.8158	
	(1.05)	(4.63)	(4.37)		(7.21)	(5.96)	
Union $density_t$	.0002	.0033	.0055	.0037	-.0011	.0009	.0051
	(0.23)	(1.47)	(2.15)	(1.36)	(0.43)	(0.27)	(1.48)
Log annual pay_t				.1865			2.5194
				(2.01)			(5.83)
Constant	.8938	.7313	1.2665	-1.9235	-2.5482	-5.3882	-24.7589
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry controls (44)	No	No	Yes	Yes	No	Yes	Yes
State fixed effects	No	Yes	Yes	Yes	Yes	Yes	Yes
2SGLS	No	No	No	No	Yes	Yes	Yes
N	1172	1172	1172	1071	1121	1121	1071
Within R^2					.8377	.8341	.7739
Between R^2					.5677	.3289	.3056
Total R^2	.8907	.9042	.9150	.9119	.7090	.5467	.5209

Notes: both pay variables are instrumented by the average weekly benefit rate. Annual wage used relates to full-time/full-year workers. All equations also include age and % male as controls. t-statistics are in parentheses.

Table 8. Happiness Equations and Likelihood-of-Finding-a-Job Equations: USA, 1973-2002 (ordered logits)

	Happiness		*Finding a Job*
Age	.0107	-.0078	-.0306
	(2.14)	(1.01)	(3.40)
Age squared	-.0001	.0002	.0001
	(0.94)	(2.24)	(0.49)
Male	-.0156	-.0454	.0045
	(0.66)	(1.89)	(0.15)
Part-time	-.0552	-.1007	.3363
	(1.47)	(2.61)	(7.32)
Temporarily not working	-.2441	-.2798	-.0433
	(3.50)	(3.82)	(0.59)
Black	-.6356	-.6399	-.2223
	(19.45)	(14.77)	(4.12)
Other non-white	-.1096	-.0984	-.1184
	(2.20)	(1.64)	(1.64)
Years of Schooling	.0585	.0538	.0759
	(13.14)	(11.49)	(11.73)
Log state unemployment	-.1427	-.1754	-.7770
	(1.78)	(2.19)	(4.79)
Workers only	No	Yes	Yes
Cut_1	-1.6354	-2.3223	-1.8992
Cut_2	1.2110	.7193	-.4370
Pseudo R^2	.0209	.0152	.0353
Log pseudo likelihood	-36893.9	-22092.6	-14678.8
N	39,998	24,725	14,032

Notes: The probability-of-finding-a-job variable is available for years 1977, 1978, 1982, 1983, 1985, 1986, 1989-1991, 1993, 1994, 1996, 1998, 2000 and 2002. All equations include 44 state dummies and also year dummies (21 for happiness and 15 for job-finding). Happiness equations also include 4 marital status dummies, and a dummy for parents divorced when respondent age 16, as in Blanchflower and Oswald (2004). Column 1 also includes dummies for unemployment, housework, student and retired. t-statistics, in parentheses, are adjusted for state level clustering. Excluded category: full-time and white.

Source: General Social Surveys

Appendix Table 1a. Average weekly benefits by state, 2000

	Average Weekly Benefit	% average weekly wage
Alabama	$159.41	.290
Alaska	$189.86	.288
Arizona	$162.51	.261
Arkansas	$210.08	.420
California	$160.00	.203
Colorado	$255.86	.360
Connecticut	$257.56	.294
Delaware	$214.85	.306
District of Columbia	$241.03	.252
Florida	$220.21	.378
Georgia	$211.89	.324
Hawaii	$283.67	.496
Idaho	$209.46	.398
Illinois	$251.58	.345
Indiana	$222.19	.374
Iowa	$238.42	.447
Kansas	$247.09	.442
Kentucky	$224.78	.409
Louisiana	$182.06	.343
Maine	$202.29	.386
Maryland	$212.51	.312
Massachusetts	$293.45	.344
Michigan	$244.12	.344
Minnesota	$290.51	.428
Mississippi	$156.62	.328
Missouri	$186.22	.311
Montana	$187.92	.414
Nebraska	$188.00	.356
Nevada	$222.43	.361
New Hampshire	$217.21	.327
New Jersey	$289.61	.345
New Mexico	$180.43	.349
New York	$247.48	.286
North Carolina	$231.21	.389
North Dakota	$210.01	.448
Ohio	$236.40	.381
Oklahoma	$214.40	.422
Oregon	$232.62	.372
Pennsylvania	$264.76	.407
Puerto Rico	$103.91	.293
Rhode Island	$253.48	.409

South Carolina	$190.18	.354
South Dakota	$180.86	.386
Tennessee	$188.74	.325
Texas	$227.11	.340
Utah	$213.89	.386
Vermont	$215.55	.391
Virginia	$203.88	.308
Virgin Islands	$198.07	.379
Washington	$280.94	.396
West Virginia	$197.53	.390
Wisconsin	$233.11	.396
Wyoming	$207.10	.408

Source: http://www.ows.doleta.gov/dmstree/uipl/uipl2k3/uipl_1703a1.htm

Appendix Table 2a. US State Log Hourly Wage Equations with Aggregate Wage Included or Average Benefits Excluded, 1980-2001

	(1)	(2)	(3)	(4)
Log hourly pay$_{t-1}$	.6806	.7156	.7257	.7473
	(34.33)	(33.72)	(36.53)	(36.34)
Log unemployment rate$_t$	-.0206	-.0441	-.0224	-.0429
	(6.15)	(8.21)	(5.93)	(7.74)
Average weekly benefit$_t$	.0002	.0003		
	(4.31)	(5.16)		
Union density$_t$	.0014	.0018	.0010	.0012
	(3.23)	(4.00)	(2.39)	(2.87)
National wage$_t$	.2683	.2152		
	(11.44)	(8.38)		
Constant	-.7890	-.4876	.0904	.2413
Year dummies	No	No	Yes	Yes
Industry controls	Yes	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes
2SGLS	No	Yes	No	Yes
State time trends	No	No	No	No
N	1122	1122	1122	1122
Within R^2		.9951		.9956
Between R^2		.9655		.9614
Total R^2	.9961	.9904	.9964	.9897
F-statistic	4113.6		3244.2	
Wald/Chi^2		1.61e+07		1.76e+07
Long run unemp. elasticity	-.0645	-.1551	-.0816	-.1698

Notes: controls are as in Table 1. National wage is defined as the average hourly wage of private sector production workers in December of each year. Source Bureau of Labor Statistics (www.bls.gov)

Appendix Table 2b. US State Log Hourly Wage Equations Using 2 Stage Procedure and Long-Run Average Unemployment, 1980-2001

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Log hourly pay$_{t-1}$	.6990	.7083	.7437	.8309	.8369	.8394	.8392
	(33.61)	(34.37)	(37.92)	(53.46)	(54.97)	(59.09)	(58.78)
Log unemployment rate$_t$	-.0246	-.0244	-.0285	-.0349	-.0349	-.0273	-.0246
	(6.50)	(6.49)	(8.58)	(12.21)	(12.24)	(7.59)	(7.46)
Average weekly benefit$_t$	.0002	.0002	.0001	.0002	.0002	.0001	.0001
	(4.03)	(3.66)	(3.14)	(4.39)	(4.48)	(3.85)	(3.96)
Union density$_t$	.0013	.0012	.0010	.0009	.0009	.0009	.0009
	(2.84)	(2.94)	(2.64)	(3.00)	(2.72)	(4.74)	(4.64)
Average Log unempt.						+.0170	
						(2.68)	
19yr Moving avge Log unempt.							+.0103
							(2.02)
Constant	.0719	-.0307	.1504	.3934	.0377	-.1034	-.1708
Including imputations	No	Yes	No	No	No	No	No
2 stage procedure	No	No	Yes	Yes	Yes	No	No
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry controls	Yes	Yes	Yes	No	No	Yes	Yes
Personal averages	Yes	Yes	Yes	Yes	No	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes	Yes	No	No
2SGLS	No	No	No	No	No	No	No
N	1122	1122	1122	1122	1122	1122	1122
R^2	.9911	.9964	.9834	.9806	.9817	.9961	.9961
F-statistic	2278.63	3179.8	174.3	433.4	479.8	3590.4	3579.9
Long run unempt. elasticity	-.0817	-.0836	-.1112	-.2064	-.2140	-.1700	-.1530

Notes: controls are as in Table 1. Column 1 is from Table 1 column 2. Two-stage procedure involves running a first stage regression and extracting the mean wage residual (coefficient on the state dummy variable). Controls at the first stage for 1979-1991 are age and its square, gender, years of education, 4 race dummies and 45 industry dummies. After 1991, years of education is replaced with 16 highest qualification dummies and 5 race dummies. Average log unemployment is the average for each state from 1979-2003.

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