

Do Greasy Wheels Curb Inequality?

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Abstract

I document a disparity in the cyclical behavior of the allocative wage across levels of educational attainment. Specifically, workers with college or more exhibit an allocative wage that is highly pro-cyclical and responsive to monetary policy shocks while high school dropouts exhibit no cyclical patterns. Meanwhile, the less educated respond to monetary policy shocks on the employment margin. Embedding these findings in an otherwise standard New Keynesian model demonstrates that heterogeneous wage rigidity results in cyclical welfare losses that exceed those of the output-gap equivalent representative agent economy. The excess welfare loss is borne by the least educated.

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1 Introduction

Heterogeneous response to monetary policy has been relatively little studied in the academic literature but receives increasingly greater attention in the popular media, particularly since the 2007 Global Financial Crisis. In this paper, I document that workers with a bachelors degree or more exhibit an allocative wage—the labor costs considered when deciding to form or dissolve an employment relationship—that is highly procyclical and sensitive to monetary policy shocks. Meanwhile high school dropouts’ allocative wage exhibits no discernible cyclical pattern. The opposite patterns across education appear in employment data: high school dropouts’ employment is more cyclically sensitive and more sensitive to monetary policy shocks. An important takeaway is that conventional monetary policy easing reduces employment inequality at the expense of raising wage inequality. I embed these findings in a simple New Keynesian model that includes price and heterogeneous wage rigidity and show that heterogeneity results in welfare losses due to fluctuations that exceed those of the output-gap and price-level equivalent representative agent economy. The excess welfare loss is borne by the least educated.

The allocative wage is notoriously difficult to measure from a macroeconomic perspective, as wage payments made today—remitted wages—are contaminated by selection effects over the business cycle and by the possibility of forward looking intertemporal smoothing imbedded in implicit labor contracts.¹ In this paper, I note that from a macroeconomic perspective the most appropriate measure of allocative wage is the analogue to a user cost as laid out in [Kudlyak \(2014\)](#) and [Basu and House \(2016\)](#). The user cost of labor (UCL) can be written as:

$$UCL_t = \mathbb{E}_t [PDV_t - \beta(1 - s)PDV_{t+1}] \quad (1.1)$$

where β is the discount factor, s the exogenous separation rate, PDV_t is the present discounted value of wage payments in an employment relationship starting at date t , and \mathbb{E}_t is the time- t expectations operator. Such a measure is required in any context in which long-term labor contracting shifts remittances intertemporally—for example, in the context of implicit labor contracts or of downward nominal wage rigidity. In any such context, remitted wages understate the degree of pro-cyclicality in the expected cost of an employment relationship. Further, the methodology for measuring the UCL , which is laid out in [Kudlyak \(2014\)](#) and employed here, addresses cyclical selection effects.

[Kudlyak \(2014\)](#) and [Basu and House \(2016\)](#) document substantial pro-cyclicality in the UCL of the representative agent. Each interprets this finding as an argument against

¹See [Daly and Hobijn \(2016\)](#) for a fresh look at selection affects and [Beaudry and DiNardo \(1991\)](#) for the canonical result regarding implicit contracts.

the common practice of assuming nominal rigidities in order to generate amplification in a Diamond-Mortensen-Pissaredes model (Kudlyak) or a New Keynesian model (Basu and House). However, this paper shows that the representative agent approach masks economically significant and policy relevant variation across educational attainment.

To this end, I re-document the well known fact that more educated workers enjoy longer term employment relationships on average. This implies that these workers are more disposed to cyclical sensitivity in the *UCL*. To see this, it is useful to manipulate the basic expression for the *UCL* to obtain:

$$UCL_t = \underbrace{w_{t,t}}_{\text{New Hire's Wage}} + \underbrace{\mathbb{E}_t \sum_{j=1}^{\infty} [\beta^j (1-s)^j (w_{t+j,t} - w_{t+j,t+1})]}_{\text{Expected Wage Wedge}}, \quad (1.2)$$

where $w_{t,t+j}$ is the wage paid at date $t+j$ to a worker hired on date t . This formulation of the *UCL* illustrates its decomposition into the *New Hire's Wage (NHW)* and the *Expected Wage Wedge (EWW)*. These two components of the *UCL* have been independently demonstrated to be pro-cyclical: [Bils \(1985\)](#) and [Beaudry and DiNardo \(1991\)](#) provide the canonical analyses, respectively. Further, this formulation highlights the key issue with the representative agent framework that motivates this paper: more educated workers, who exhibit lower separation rates, are differentially exposed to the *EWW*.

Indeed, I document three important facts. First, highly educated workers experience substantial pro-cyclicalities in their *UCL*, while their less-educated counterparts experience a-cyclicalities. Second, this result stems from differential exposure to the *EWW* and is amplified by differential sensitivity of the *EWW* to cyclical conditions for the more educated. Third, the sensitivity of the representative agent's *UCL* to cyclical indicators stems largely from the procyclicality of the *UCL* of the most educated. This documented heterogeneity across education revives the possibility of nominal rigidities as a source of amplification in a model with a richer set of agents and raises the concern that, in such an economy, monetary policy may induce inequality.

I address this concern empirically and document that monetary policy contractions robustly decrease the *UCL* of the most highly educated while having limited effects on the *UCL* of the less educated. This interacts with the steady-state inequality across groups, generated by well-documented returns to education, and leads to attenuation of inequality as measured by wages in response to monetary policy contractions.² With respect to

²The baseline specification considers the years between Chairman Volcker's appointment in 1979 and the onset of the financial crisis in 2007 (excluding 2007) and monetary policy shocks identified using the method of [Romer and Romer \(2004\)](#) with Greenbook forecasts updated as in [Coibion \(2012\)](#). Results are robust to excluding the Volcker Reform (1979-1982).

employment, monetary policy contractions have the opposite effect. My baseline specification reveals that monetary policy contractions decrease employment rates among the less educated while having no effect on the most educated.³ Again, this interacts with the steady-state employment inequality across groups and leads to amplification of employment inequality.

I embed these facts in a standard New Keynesian model featuring both price and wage rigidities based on [Erceg et al. \(2000\)](#). My key departure is to allow wage rigidities to be differentially binding across “varieties” of workers, where varieties are meant to capture the documented differences in wage and labor elasticities with respect to monetary policy shocks. I assume that, up to heterogeneous wage rigidities, worker varieties are identical.⁴ Meanwhile I allow income pooling only within variety. In this framework, I show three things. First, the elasticity of earnings and consumption to an aggregate demand shock are identical across varieties. Second, the price and output elasticities with respect to the interest rate shock are identical to a homogeneous agent model in which the wage elasticity is a linear combination of the heterogeneous wage elasticities where the linear combination is determined by the output elasticities of the labor varieties. Third, the elasticity of the analogously constructed linear combination of period utility in the heterogeneous agent economy is larger than the elasticity of period utility in the representative agent economy, with the excess welfare losses being born by the variety with less flexible wages (the less educated).

The modeled environment lends itself to quantitative welfare analysis in the framework of [Galí et al. \(2007\)](#). Under a baseline calibration with both Frisch elasticity and elasticity of intertemporal substitution (EIS) equal to unity for all varieties, I find that the welfare loss due to cyclical inefficiencies in the heterogeneous agent economy exceeds that of the output-gap equivalent representative agent economy by more than fifteen percent.⁵ Within the heterogeneous agent economy itself, the welfare loss experienced by those with less than a high school education exceeds that of those with at least a college degree more substantially: by more than 15 times in the baseline calibration!

This paper fits into several literatures. Most directly, I contribute to answering questions regarding the heterogeneous effects of monetary policy ([Romer and Romer, 1999](#); [Coibion](#)

³Similar to the facts documented in [Ramey \(2016\)](#), results regarding employment are sensitive to the exclusion of the Volker Reform. When these years are excluded, monetary policy has no statistically discernable impact on employment for any level of educational attainment.

⁴All results extend trivially to the realistic addition to the model of differential but log-separable labor efficiency across varieties.

⁵Note, variation in the sensitivity of wages to cyclical position and monetary policy shocks could also be indicative of differential Frisch elasticity across education. However, I document empirically that a substantial portion of heterogeneity in the sensitivity of wages derives from sensitivity of the EWW. This suggests binding constraints on adjusting wages in the near term that are differentially easy to smooth through for more highly educated workers with longer expected employment relationships.

et al., 2017; White, 2018; Cravino et al., 2018; Aaronson et al., 2019). As noted, my findings suggest that monetary policy easing, at least using conventional instruments, decreases employment inequality but increases wage inequality. This suggests consideration of the cyclical variation in inequality (at least on these margins) and welfare that the monetary policy maker need tolerate if she exclusively aims to minimize the output gap and maintain stable prices. Excess welfare loss relative to the representative agent representation, which are driven by losses experienced by the less educated, suggests that policy makers could improve their policy rule by also targeting a weighted unemployment gap that places relatively more weight on the unemployment of the less educated.

This paper also contributes to the literature on monetary policy in the presence of durables, for example, Barsky et al. (2007). In the case of a durable and nondurable capital good, intertemporal substitution between investment in the inputs reduces the economy's sensitivity to monetary policy shocks. The empirical facts documented in this paper suggest that similar mechanisms are at work in the labor market—greater expected durability of the employment relationship with a highly educated worker contributes to greater flexibility of the associated allocative wage—with the negative consequence of cyclical variation in inequality and welfare. In addition, the output-gap and price level equivalent representative agent's wage elasticity is a weighted average of the cyclical wage elasticity of the underlying agents.

This in turn relates to the literature on cyclical sorting and the cleansing versus sullyng effects of recessions.⁶ The results here suggest that variation in labor market composition over the business cycle is both a consequence of variation in match quality generated through variation in frictions in labor market flows and of cost minimization that takes into account that differential wage rigidities lead to cyclical variation in wage differentials, documented here. The evidence documented here suggests that the distortionary effects of differential rigidities are substantial relative to frictions in labor market flows.⁷ Thus, in addition to the welfare costs of the sullyng effect of recessions on output via match quality, there is additional welfare loss due to the distortion of relative wages.

Lastly, these results contribute to the literature on the nature of the labor contract. Often it is assumed that a joint surplus that is increasing in educational attainment is needed to justify greater employment stability among the highly educated. These results, however, suggest that instead greater wage flexibility may drive the differences and that *flow* joint surpluses need not vary substantially with education. An important extension to consider

⁶A noncomprehensive list of highlights include Barlevy (2002), Kahn (2010), Hagedorn and Manovskii (2013), Kahn and McEntarfer (2014), Haltiwanger et al. (2015), Cairó et al. (2016), Abel and Deitz (2016), Haltiwanger et al. (2018), and Crane et al. (2018).

⁷In particular, results are robust to controlling for cyclically in match quality using the method of Hagedorn and Manovskii (2013).

is the impact of differential durability in a microfounded wage contract à la [Elsby \(2009\)](#) or [Rudanko \(2009\)](#). Finally, my results are corroborated by differential flexibility in near-term measures of wages—for example, differential use of bonuses for the highly educated ([Grigsby et al., 2018](#)).

The remainder of this paper proceeds as follows. Section 2 describes the data used. Section 3 documents variation in the stability of employment across educational attainment. Section 4 describes the empirical methodology and documents the aforementioned facts regarding the cyclicity of the *UCL* and its components. Section 5 documents the heterogeneous reaction of employment and *UCL* to monetary policy shocks across educational attainments. Section 6 details parsimonious assumptions under which documented wage and employment sensitivities imply more sensitivity of welfare for the less skilled and measures the welfare costs of heterogeneity using the framework of [Galí et al. \(2007\)](#). Section 7 concludes. The Appendix provides further documentation of the data and documents robustness to alternative detrending methods, treatments of educational upgrading on the job, and identification methods of monetary policy shocks. I also show that differential sensitivity to monetary policy shocks is robustness to the inclusion of a stochastic discount factor and time-varying separation rates.

2 Data

Data used in this paper come primarily from two sources.

National Longitudinal Survey of Youth 1979

The National Longitudinal Survey of Youth (NLSY) contains panel data on employment histories and wages. The data are an unbalanced panel of workers surveyed yearly from 1979 to 1994 and every other year thereafter. The initial sample contains 12,686 respondents who were between 14 and 21 years old at the date of the intimal survey. Although the sample is not representative of the U.S. population, yearly cross-sectional sampling weights render the sample comparable to each years population up to the natural aging of the sample. Following [Kudlyak \(2014\)](#) and [Basu and House \(2016\)](#), I restrict the sample to males.

Retrospective wage and employment date information is available for each respondent in each survey for up to five jobs. From these data, the NLSY constructs a variable “hourly rate of pay” to synchronize reporting pay intervals (hour, day, week, month, year) using reported typical hours worked. This variable includes tips, overtime pay, and bonuses before any deductions. Following [Basu and House \(2016\)](#) I deflate the hourly rate of pay with the implicit price deflator for the non-farm business sector.

Current Population Survey

The Current Population Survey (CPS) records the employment status and demographic characteristics of a representative sample of U.S. workers monthly. I use the data since 1976 over which time micro data are available in the basic monthly survey that enable consistent construction of unemployment and employment rates by educational attainment. From 1982 onward the micro data collected from the outgoing rotation groups—roughly one-fourth of the sample—also contains information on usual hours worked.

Educational attainment is also recorded in the CPS. Prior to 1992, education is recorded as the highest grade of school completed. From 1992 onward, education is recorded as the highest degree or diploma attained. The difference is particularly important for consistently measuring high school graduation. I follow the crosswalk used in [Elsby and Shapiro \(2012\)](#). Details of the crosswalk and unemployment rate series by education implied by the crosswalk are in the Appendix.

Other data

I supplement these main data sources with data on labor market tightness constructed as in [Barnichon \(2010\)](#) using the publicly available data from the Conference Board and JOLTS. I also construct monetary policy shocks as in [Romer and Romer \(2004\)](#), using data updated by [Coibion \(2012\)](#). In assessing the welfare implications of heterogeneity, I use the following series drawn from the USECON database: compensation per hour (LXNFC) and real and nominal output (LXNFO and LXNFI), which refer to the nonfarm business sector; nondurable and services consumption (CNH + GSH), drawn from the respective NIPA series; and implicit price deflator (LXNFI).

In robustness tests I consider monetary policy shocks as identified by [Gertler and Karadi \(2015\)](#) and discount rate shocks as identified by [Hall \(2017\)](#) using data obtained from the respective replication files. I also consider variation in the separation rate using a linkage of the CPS Monthly survey survey based on replication code from [Shimer \(2012\)](#).

3 Employment Inequality

I begin by documenting some well-known facts regarding employment inequality in the CPS. The first two columns of table 1 report the mean and standard deviation of the aggregate unemployment rate as well as the unemployment rate for workers with less than high school, high school or some college, and college or more. The third column presents the standard deviation of each series after detrending using the method of [Hamilton \(2018\)](#). As is well

Table 1: Labor Market Volatility by Education (25+ years).

	Employment Rate ^a			Hours per Week ^b		
	Mean	Standard Deviation	Volatility	Mean	Standard Deviation	Volatility
All	6.61	1.63	1.34	38.58	0.58	0.25
< High School	12.84	2.67	2.12	34.58	0.78	0.39
High Sch. / Some Coll.	6.25	1.74	1.44	38.30	0.51	0.30
≥ Bachelors	2.90	0.79	0.66	41.06	0.53	0.28

^a *Source:* Current Population Survey Monthly Data 1976-2018 and author’s calculations.

^b *Source:* Current Population Survey Outgoing Rotations 1982-2018 and author’s calculations. Hours per week are conditional on employment.

Note: Detrended unemployment rate reports the standard deviation of the data expressed in terms of deviations from trend. Series are detrended separately by education. Trend is recovered using the method of [Hamilton \(2018\)](#). All data are at the quarterly frequency.

known, unemployment rates are higher and volatility greater for workers with less education.⁸ The differences are significant, with high school dropouts experiencing more than three times the volatility experienced by college-educated workers. A similar pattern is observable for hours. More highly educated workers work more, and less-volatile hours.⁹

Employment volatility, naturally, is inversely related to the durability of labor contracts. Data from the NLSY reveal that, indeed, completed tenure and job cycles—defined as the time between consecutive unemployment spells—are longer for workers with greater educational attainment. These are reported in the first two columns of table 2. The differences are again significant. College educated workers enjoy tenures more than twice as long and job cycles more than three times as long as high school-educated workers. The final column documents yearly separation rates by educational attainment. The pattern persists, with more-educated workers enjoying a lower probability of job separation.

4 Cyclicalities of Allocative Wages by Education

The “wage” is a notoriously difficult macroeconomic object to measure. Not only is there substantial quantitative divergence between the various measures put forth – average hourly earnings, new hires wages, etc. – but there is disagreement about which measure is substantively correct. In this paper, I argue that the appropriate measure of allocative wage to

⁸Results are in line with [Cairó and Cajner \(2018\)](#), differing only due to slightly different time horizons and the use of the [Hamilton \(2018\)](#) filter as opposed to the HP filter.

⁹As I discuss in the following section, the *UCL* appropriately measures the allocative wage in the absence of financial frictions. Consequently, the remainder of the paper focuses on data up until the Global Financial Crisis. In Appendix A.1, I replicate table 1 in the truncated data. The takeaways are the same.

Table 2: Tenure, Job Cycle Length, and Separation Rates by Education.

	Duration (years) of		Separation Rate (yearly)
	Tenure	Job Cycle	
All	2.44	4.97	0.29
< High School	1.41	2.55	0.38
High School / Some College	2.44	4.92	0.29
≥ Bachelors	3.70	8.17	0.24

Source: National Longitudinal Survey of Youth 1979 and author’s calculations.

consider from the macroeconomic perspective is the user cost of labor as defined by Kudlyak (2014). This measure takes into account, but does not impose, the possibility that labor market frictions impart a durable quality to an employment relationship and that as a result the sequence of payments under a(n implicit) wage contract might diverge from the sequence of wages that would arise in a spot market. Importantly, given the above analysis, one should expect that any such divergence would be more extreme for employment relationships that are more durable and therefore among the most educated. Further, the methodology proposed by Kudlyak (2014) also addresses issues of composition bias that plague many other macroeconomic wage measures.

4.1 Measuring the User Cost of Labor

The first step in measuring the user cost of labor is to clean the wage data from individual-specific effects using an augmented Mincer regression which includes, among other things, individual fixed effects. I follow the basic empirical specification of Kudlyak (2014) and Basu and House (2016):

$$\ln w_{t,\tau,E}^i = c_E + \alpha_E^i + \zeta_E t + \Phi_E X_{t,E}^i + \sum_{d_0=1}^T \sum_{d=d_0}^T \chi_{d_0,d,E} D_{d_0,d}^i + \varepsilon_{t,E}^i \quad (4.1)$$

Here, $w_{t,\tau,E}^i$ is the real wage for individual i at time t who was hired at time τ and has education $E \in \{\textit{less than high school, high school/some college, college or more}\}$. This regression provides a best linear prediction of the log real wage at time t of a worker i who started a job in period τ and has education E . In its most general form, this wage regression allows for a time trend, demographic and industry controls (included in $X_{t,E}^i$), individual fixed effects (the α_E^i coefficients), and time effects that depend on two periods: when the individual began work at the current job and the current date. The additional covariates in

the $X_{t,E}^i$ matrix are the individual's experience at time t (and experience squared), tenure at time t (and tenure squared), schooling completed, and industry fixed effects. Experience is defined as the maximum of ($Age - 6 - years\ of\ schooling$) and 0. The dummy variables $D_{d_0,d}^i$ take the value 1 if $d_0 = \tau$ and $d = t$ and 0 otherwise.

I aim to construct the user cost of labor separately by educational attainment. To this end, I split the NLSY79 sample into three sub-samples based on the educational attainment of the respondent at the time of hiring. Splitting the sample ex-ante allows all the coefficients in the augmented Mincer regression to differ by educational attainment.¹⁰

The χ_E coefficients in regression equation 4.1 are particularly important for constructing the user cost of labor. At time t , all education E workers who began work at their current job at date- τ get an additional adjustment to their predicted wage given by the coefficient $\chi_{\tau,t,E}$. These adjustments imply that workers who begin at date- τ experience an expected strip of log wage realizations given by $\{\widehat{\chi}_{\tau,\tau,E}, \widehat{\chi}_{\tau+1,\tau,E}, \widehat{\chi}_{\tau+2,\tau,E}, \dots, \widehat{\chi}_{\tau+j,\tau,E}, \dots\}$ etc. These dummy variables thus adjust for vintages of hired workers, where the vintage is defined by when the worker was hired in addition to the current calendar date. Notice that the variable $\widehat{\chi}_{\tau,\tau,E}$ reflects the wages of a newly hired worker (i.e., the date- τ wage of a worker hired at date- τ). Following Basu and House (2016), I truncate the χ strips at seven years (including year 0).¹¹

I base my calculation of the user cost of labor on equation (1.1):

$$\widehat{PDV}_{t,E} = \sum_{j=0}^{\infty} \beta^j (1 - s_E)^j \exp \left\{ \widehat{\ln} w_{t,t+j,E} \right\}. \quad (4.2)$$

Note that in addition to requiring a sequence of predicted log wages $\{\widehat{\ln} w_{t,t+j,E}\}_{j=0}^{\infty}$, this calculation requires a separation rate s_E and a discount factor β . I estimate s_E as the average annual separation rate by education group, which suggests setting s equal to 0.38, 0.31 and 0.25 for less than high school, high school/some college, and college or more respectively. I assume the annual discount factor is set to 0.97. Note that my calculation of the present value of wage payments is truncated at seven years (including the initial year).¹²

¹⁰Splitting the sample ex-ante could spuriously increase the cyclically of the measured allocative wage for less-educated workers if educational upgrading on the job is more likely to occur for spells initiated in expansions. In the Appendix, I document the rate of educational upgrading by decade and by cyclical position. I do not find strong evidence suggesting bias. Further, results are robust to splitting the sample on contemporaneous education and on highest educational attainment achieved during the job spell. The former effectively treats an educational upgrade as a new hire, and the latter assumes that employers have ex-ante knowledge of an employees education prospects at the time of hiring.

¹¹More precisely, I include all of the dummy variables in the estimation of (4.1); however, following Kudlyak (2014) and Basu and House (2016), I use only seven χ estimates when I calculate the user cost of labor.

¹²This may attenuate the differences between the allocative wage across educational attainment because

To construct the projected wage payments $\widehat{\ln w}_{t,\tau}$, I consider the anticipated wage payments for a firm that hires an average worker at date- t from each education group. As the employment relationship continues, the measure of the worker’s experience and the measure of the workers tenure both increase. I assume that the initial experience is fixed at the sample average for each education group: 11.29, 13.28, and 12.33 years for less than high school, high school, some college, and college or more, respectively. I set the initial tenure variable to 0.5 years (which implicitly assumes that a worker who reports being newly hired at his current job at the time of the interview was hired 6 months earlier). Average years of schooling are 9.8, 12.6, and 16.8 years for less than high school; high school or some college; and college or more, respectively. Based on (4.1), at date- τ , a worker hired at date- $t \leq \tau$ has a projected log wage

$$\widehat{\ln w}_{t,\tau} = \widehat{c} + \widehat{\zeta}\tau + \widehat{\Phi}\bar{X}_{\tau-t} + \widehat{\chi}_{\tau,t}, \quad (4.3)$$

where $\bar{X}_{\tau-t}$ are demographic controls for the average worker within each education group.

4.2 Cyclically by Education Group

Here I test the cyclical properties of the allocative wage along-side more traditional wage measures: the new hire’s wage and average hourly earnings (adjusted for observable). New hires’ wage is measured simply as

$$\widehat{\ln w}_{t,t} = \widehat{c} + \widehat{\zeta}t + \widehat{\Phi}\bar{X} + \widehat{\chi}_{t,t} \quad (4.4)$$

with all coefficients estimated as in equation 4.1 above.

I estimate average hourly earnings from the NLSY data by estimating

$$\ln w_{t,E}^i = c_E + \alpha_E^i + \Phi_E X_{t,E}^i + \sum_{d_0=1}^T \omega_{d,E} D_d^i + \varepsilon_{t,E}^i, \quad (4.5)$$

where, as before, $X_{t,E}^i$, contains educational attainment, a quadratic in potential experience and tenure, and industry fixed effects.

Following [Basu and House \(2016\)](#), I check the cyclical properties of each wage measure by regressing wage on either detrended log real GDP or detrended national unemployment

highly educated workers are more likely to have truncated spells.

Table 3: Wage Cyclically by Education Group (1978 - 2006).

Cyclical Indicator = log real GDP ^a	User Cost of Labor		New Hire's Wage		Avg. Hourly Earnings ^b	
< High School	-0.26	(0.58)	-0.31	(0.25)	-0.26**	(0.14)
High School / Some Coll.	0.95*	(0.55)	-0.03	(0.27)	0.01	(0.21)
≥ College	3.02*	(1.53)	1.28**	(0.49)	0.25	(0.31)
Observations	29		29		29	

Cyclical Indicator = unemployment rate ^a	User Cost of Labor		New Hire's Wage		Avg. Hourly Earnings ^b	
< High School	0.06	(1.86)	-0.56	(0.82)	-0.15	(0.48)
High School / Some Coll.	-6.22***	(1.33)	-1.39*	(0.81)	-1.29**	(0.61)
≥ College	-9.31*	(4.68)	-6.33***	(1.14)	-2.46***	(0.80)
Observations	29		29		29	

Note: All regressions control for a quadratic time trend. Standard errors in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

Source: National Longitudinal Study of Youth 1979 and author's calculations.

^a Detrended using the filter proposed by Hamilton (2018).

^b Controlling for experience, industry fixed effects, and individual fixed effects.

rate and a quadratic in time.¹³

$$\log(\text{Wage Measure}) = \text{constant} + \beta * \text{Cyclical Indicator} + \gamma_1 * \text{year} + \gamma_2 * \text{year}^2 + \varepsilon$$

Regressions are weighted according to the size of the contemporaneous NLSY sample.

Table 3 reports the coefficients from these regressions. Strikingly, the *UCL* exhibits no cyclicity for the least educated but is increasingly procyclical as educational attainment increases. This pattern holds more weakly for new hire's wages and even more weakly for average hourly earnings after controlling for compositional biases.

Discussion

The pattern of increasing procyclicality in educational attainment is notable for two main reasons.

¹³To avoid the endpoint problem in HP filtering, I detrend the data using the filter recently proposed in Hamilton (2018), which does not suffer from the endpoint problem. In the Appendix, I show that results are robust to following the approach of Basu and House (2016), who address the endpoint problem by adding 120 months of predicted unemployment rates taken from an estimated AR(6) to the end of the sample and then HP filtering the padded monthly series using a smoothing parameter of 500,000.

First, wages of the least educated are a-cyclical by all measures. A-cyclical for this group reverses the headline finding of [Kudlyak \(2014\)](#) and [Basu and House \(2016\)](#) for this segment of the labor market. Further, this a-cyclical potentially restores the potential of nominal wage rigidity in generating both amplification and persistence in the Diamond-Mortensen-Pisaredes class of models, criticized by [Kudlyak \(2014\)](#), and in the class of New Keynesian models, criticised by [Basu and House \(2016\)](#).

Second, increasing pro-cyclical as educational attainment rises suggests that jobs of more educated workers are more robust to changing business cycle conditions. This finding suggests that policy may have a differential role in mitigating shocks for each group. I document evidence of differential impacts of monetary policy shocks in [Section 5](#).

Before turning to the interaction between monetary policy and the differential wage rigidity that is suggested by this evidence, I document that my findings are robust to the key criticism of the [Beaudry and DiNardo \(1991\)](#) evidence for implicit contracts—upon which this work builds—leveled by [Hagedorn and Manovskii \(2013\)](#). I then discuss the origins of differing cyclically across education and their contribution to the result documented by [Kudlyak \(2014\)](#) and [Basu and House \(2016\)](#) for the representative agent.

4.3 Robustness to Cyclical Sorting

[Hagedorn and Manovskii \(2013\)](#) argue that the classic result of [Beaudry and DiNardo \(1991\)](#) is a spurious product of cyclical variation in the match quality of jobs. In particular, they argue that during expansions, workers have differentially greater opportunities to reallocate to more suitable jobs. [Hagedorn and Manovskii \(2013\)](#) propose proxies for match quality built on the intuition of a job ladder model: M_c is the cumulative labor market tightness experienced between the end of the most recent employment spell and the start date of the current job, and M_j is the cumulative labor market tightness experienced during the completed tenure at the current job.¹⁴ They demonstrate that inclusion of these proxies in the classic [Beaudry and DiNardo \(1991\)](#) regression eliminates key evidence of implicit contracts: the statistical significance of the lowest unemployment rate experienced during a job cycle in predicting the wage.

In the context of the present exercise, the criticism of [Hagedorn and Manovskii \(2013\)](#) suggests that the wage premium enjoyed by workers hired during expansions is simply a product of their associated differential match quality. To address this criticism, I augment

¹⁴Appendix [A.4](#) gives a detailed description of how NLSY data are mapped into job cycle and tenure. Labor market tightness is measured using the publicly available data from the Conference Board and JOLTS as in [Barnichon \(2010\)](#).

Table 4: Controlling Match Quality.

Cyclical Indicator = log real GDP ^a	User Cost of Labor		New Hire's Wage		Avg. Hourly Earnings ^b	
< High School	-0.28	(0.56)	-0.34	(0.22)	-0.23*	(0.13)
High School / Some Coll.	0.85*	(0.48)	-0.09	(0.24)	0.03	(0.18)
≥ College	2.88*	(1.43)	1.07**	(0.46)	0.27	(0.29)
Observations	29		29		29	

Cyclical Indicator = unemployment rate ^a	User Cost of Labor		New Hire's Wage		Avg. Hourly Earnings ^b	
< High School	0.12	(1.83)	-0.32	(0.75)	-0.09	(0.44)
High School / Some Coll.	-5.40***	(1.16)	-1.02	(0.74)	-1.15**	(0.52)
≥ College	-7.84*	(4.48)	-5.57***	(1.08)	-2.30***	(0.78)
Observations	29		29		29	

Note: All regressions control for a quadratic time trend. Standard errors in parentheses.

*** p<0.01, ** p<0.05, * p<0.1.

Source: National Longitudinal Study of Youth 1979 and author's calculations.

^a Detrended using the filter proposed by Hamilton (2018).

^b Controlling for experience, industry fixed effects, and individual fixed effects.

equation 4.1 with the proposed proxies for match quality:

$$\ln w_{t,\tau,E}^i = c_E + \alpha_E^i + \zeta_E t + \Phi_E X_{t,E}^i + \gamma_{c,E} M_c^i + \gamma_{j,E} M_j^i + \sum_{d_0=1}^T \sum_{d=d_0}^T \chi_{d_0,d,E} D_{d_0,d}^i + \varepsilon_{t,E}^i \quad (4.6)$$

If match quality explains the cyclical variation (and is well measured by M_c and M_t) then the entire block of χ coefficients should fall to zero and the resulting UCL should be acyclical.

Table 4 reproduces table 3 using the parameter estimates from the augmented equation 4.6. Average cumulative labor market tightness during pre-job employment experience is 28.7, 58.4, and 99.1 for less than high school, high school / some college, and college or more, respectively. Average cumulative labor market tightness completed job tenure is 83.4, 144.9, and 202.6 for less than high school, high school / some college, and college or more, respectively. As expected, controlling for match quality accounts for some of the cyclicity of wages; however, the broad picture remains the same. More educated workers experience more cyclically sensitive wages, even conditional on their match quality.

Henceforth, all analysis controls for match quality.

Table 5: Augmented Mincer Regressions.

	< High School		High School / Some College		≥ College	
Tenure	0.029*** (0.003)	0.024** (0.009)	0.029*** (0.001)	0.014*** (0.003)	0.018*** (0.003)	0.010 (0.010)
(Tenure) ²	-0.001*** (0.000)	-0.001** (0.000)	-0.001*** (0.000)	-0.000*** (0.000)	-0.000 (0.000)	-0.000 (0.000)
Observations	10,525	10,381	35,899	35,499	10,400	10,279
R-squared	0.590	0.619	0.647	0.660	0.686	0.705
Spell FE		X		X		X
Match quality	X	X	X	X	X	X
P. Exp. (quadratic)	X	X	X	X	X	X
Year (linear)	X	X	X	X	X	X
Industry FE	X	X	X	X	X	X
Worker FE	X	X	X	X	X	X

Source: National Longitudinal Study of Youth 1979 and author's calculations. The dependent variable is the natural logarithm of real hourly wage. All regressions are estimated with fixed effects using sampling weights. Unemployment rate is the annual unemployment rate.

Robust standard errors in parentheses.

*** p<0.01, ** p<0.05, * p<0.1.

4.4 Differential Separation Rates or Wage-Tenure Effects?

I now turn to decomposition of the established heterogeneity in wage cyclicality. Intuition, based on examination of equation 1.2, suggested that this derives from differentials in separation rates. The data reveal that these differentials are amplified by differentials in the cyclical sensitivity of the wage-tenure profile across educational attainment.

I begin by considering the effect of including the block of start date dummies in the Mincer-type regression. Table 5 reports the coefficient on tenure in a baseline Mincer-type regression and the regression augmented with start-date dummies, by education. From this one can see that for the most educated, a large part of the return to tenure is captured by the time during which the tenure was accrued. Meanwhile, the return to tenure for the least educated is largely invariant to the inclusion of this additional information.

This illustrates that in their representative agent exercises, both Kudlyak (2014) and Basu and House (2016) obtain spurious results for two reasons. First, and most straightforward, the *UCL* is averaged across types with differing cyclical sensitivities. Second, the coefficients χ fail to account for heterogeneity by education and are biased toward the χ experienced by the educated. The bias derives from the disproportionate representation of the educated at

Table 6: Job Duration v.s. Wage-Tenure Profiles: Cyclical Regressions.

Cyclical Indicator = log real GDP ^a	User Cost of Labor		Holding Constant:			
			Separation Rate		Wage-Tenure Effects	
< High School	-0.28	(0.56)	-0.26	(0.74)	0.47	(0.42)
High School / Some Coll.	0.85*	(0.48)	0.84*	(0.47)	0.52	(0.53)
≥ College	2.88*	(1.43)	2.36*	(1.18)	0.59	(0.65)
Observations	29		29		29	

Cyclical Indicator = unemployment rate ^a	User Cost of Labor		Holding Constant:			
			Separation Rate		Wage-Tenure Effects	
< High School	0.12	(1.83)	0.11	(2.40)	-3.48**	(1.18)
High School / Some Coll.	-5.40***	(1.16)	-5.37***	(1.15)	-3.95**	(1.51)
≥ College	-7.84*	(4.48)	-6.58*	(3.69)	-4.45**	(1.90)
Observations	29		29		29	

Note: All regressions control for a quadratic time trend. Standard errors in parentheses.

*** p<0.01, ** p<0.05, * p<0.1.

Source: National Longitudinal Study of Youth 1979 and author's calculations.

^a Detrended using the filter proposed by Hamilton (2017).

long tenures, which is itself a consequence of their differentially small separation rates.

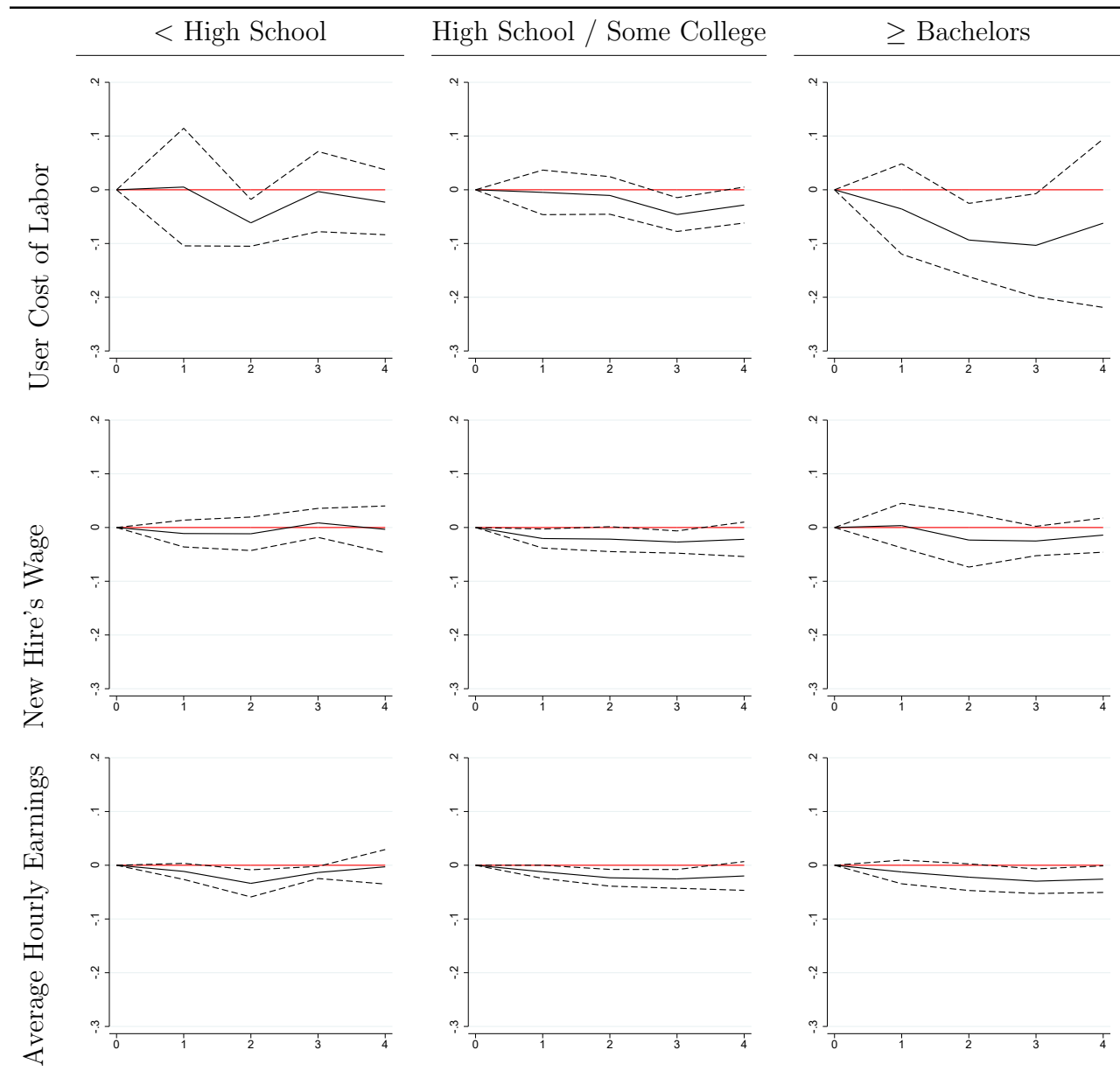
To illustrate the second point, I construct counterfactual *UCL* by education under two assumptions: (1) constant separation rates across education and (2) constant cyclical sensitivity of wage-tenure profiles across education. Table 6 reports both counterfactuals. This exercise clearly shows that differences in the cyclicity across education groups are mainly attributable to differences in the sensitivity of wage-tenure profiles to the aggregate state.¹⁵

5 Monetary Policy Shocks

In this section, I document the impulse responses of unemployment and wages to monetary policy shocks. My approach is straightforward. First, monetary policy shocks are identified from Greenbook forecast errors as in Romer and Romer (2004). Second, I trace out the impulse response of unemployment and wages by educational attainment using the local projection method of Jordà (2005).

¹⁵Following the literature, this paper considers the male NLSY79 sample. However, variation across gender is informative. In particular, the female sample exhibits both higher rates of separation and less cyclically sensitive *UCL* and *EWV* at all levels of educational attainment.

Figure 1: Impulse Response to a 100 basis point Monetary Policy Contraction: Wages by Education.



Note: 95% confidence interval.

Source: National Longitudinal Study of Youth 1979, Greenbooks as cleaned by Coibion (2012), and author's calculations.

Specifically, as in Romer and Romer (2004), I identify monetary policy shocks as the forecast error when policy is predicted using economic information in the Federal Reserve's Greenbook forecasts. Thus, a monetary policy shock at date m , ε_m is estimated as the

residual of the following regression:

$$\begin{aligned} \Delta ffb_m = & \alpha + \beta ffb_m + \sum_{i=-1}^2 \gamma_i \tilde{\Delta}y_{m,i} + \sum_{i=-1}^2 \lambda_i \left(\tilde{\Delta}y_{m,i} - \tilde{\Delta}y_{m-1,i} \right) + \sum_{i=-1}^2 \varphi_i \tilde{p}_{m,i} \\ & + \sum_{i=-1}^2 \theta_i (\tilde{\pi}_{m,i} - \tilde{\pi}_{m-1,i}) + \rho(\tilde{u})_{m,0} + \varepsilon_m \end{aligned}$$

where ffb_m is the federal funds target at the time of the meeting, $\tilde{\Delta}y_{m,i}$ denotes the forecast of real output growth, $\tilde{p}_{m,i}$ denotes the forecast of inflation, and $\tilde{u}_{m,0}$ denotes the forecast of current unemployment. The index i is the horizon of the forecast with -1 being the previous quarter. Using the data series in [Romer and Romer \(2004\)](#) extended by [Coibion \(2012\)](#), I recover shocks for the period 1969 to 2008.

Following the local projection method of [Jordà \(2005\)](#), I estimate the impulse response at horizon h to a 100 basis point contractionary monetary policy shock:

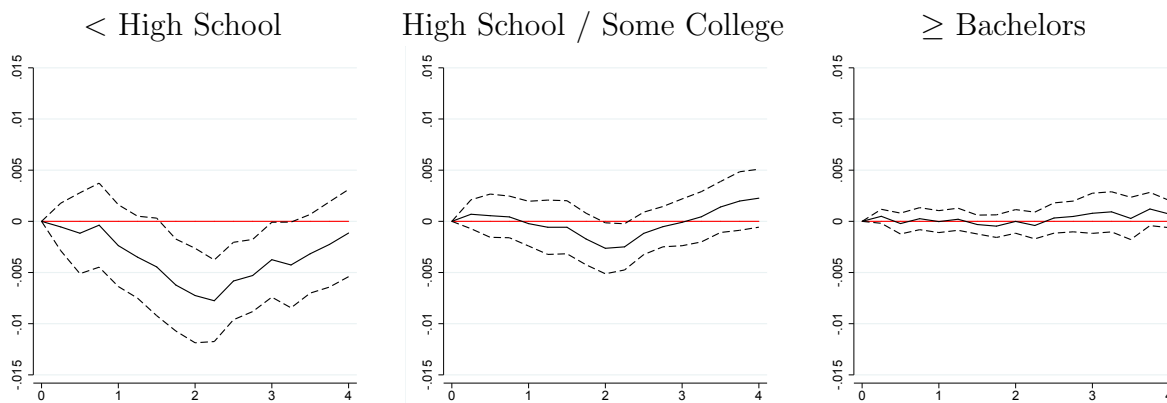
$$outcome_{t+h} = \alpha_h + monetary\ policy\ shock_t \beta_h + controls_t \gamma_h + v_{t+h}$$

where $controls_t$ contains lags of the monetary policy shock, the lags of the outcome variable, and contemporaneous and lagged unemployment rate for all workers 25 and older. I select a lag length corresponding to one year for unemployment and three years wage data. This choice reflects the notion that monetary policy makers may react to the contemporaneous remitted wage realization, which the previous section showed to be less volatile than the allocative wage. Meanwhile, deviations between the remitted and allocative wage are revealed in subsequent periods. Standard errors are as in [Newey and West \(1987\)](#). In wage regressions data are weighted according to the size of the underlying contemporaneous NLSY sample.

As in section 4, average hourly earnings, new hire's wages, and the user cost of labor are computed from the yearly and bi-yearly micro-data of the NLSY. Impulse responses are plotted in [Figure 1](#). I find that average hourly earnings and new hires' wages are largely unresponsive to monetary policy shocks. Strikingly, however, the user cost of labor falls markedly in response to a monetary policy contraction for workers with higher educational attainment. At a two- and three-year horizon, the *UCL* of college graduates falls approximately 10 basis points in response to the 100 basis point monetary policy contraction. On the other end of the spectrum, wages of high school dropouts exhibit a much more mild response.

As in section 3, employment by education are computed from the monthly microdata of the CPS. I aggregate these to the quarterly frequency. Impulse responses are plotted

Figure 2: Impulse Response to a 100 basis point Monetary Policy Contraction: Employment by Education.



Source: Current Population Survey Basic Monthly Files, Greenbooks as cleaned by Coibion (2012), and author’s calculations.

in Figure 2. As in Coibion et al. (2017), I find that employment rates fall in response to monetary policy shocks. Again, I find marked differences by education. At a two-year horizon, the employment rate of high school dropouts falls by about three-fourths of a basis point in response to the 100 basis point monetary policy contraction. On the other end of the spectrum, employment of college graduates is unaffected.¹⁶

¹⁶In the Appendix, I consider alternative strategies for the identification of monetary policy shocks. Figure A2 plots the impulse response of the employment rate, hours, and the user cost of labor when the Volker Reform (1979-1982) is excluded. As documented in Coibion (2012) and Ramey (2016), I find that the estimated response of employment is sensitive to the inclusion of the early 1980’s. This is somewhat mitigated by hours, as the second row of Figure A2 shows that hours may still be sensitive to monetary policy shocks for the least skilled even in this latter period. Finally, results regarding the *UCL* are robust to the exclusion of the Volker reform. In Figure A3, I check the robustness of the identification of the underlying shocks to inclusion of the 1969-1978 Greenbook data. Again, consistent with the literature, I find that results regarding employment are not robust. On the other hand, sensitivity of the *UCL* to these alternatively identified shocks remains.

Finally, I consider identifying monetary policy shocks using a high frequency identification strategy as in Gertler and Karadi (2015). This identification strategy is possible in the period after 1990. Figure A5 shows that using this identification strategy I identify “wrong signed” fluctuations in employment and hours and very large standard errors on predictions for the *UCL* rendering them statistically indistinguishable from zero. However, for several reasons, I place little weight on these results. As pointed out by Ramey (2016), the Gertler and Karadi (2015) shocks are (1) not mean zero, (2) autocorrelated, (3) predictable using Greenbook forecasts, and 4) provide inconsistent results when impulse responses are identified using a SVAR versus the Jordà (2005) local projection method. The finding of “wrong signed” employment comports well with the finding of “wrong signed” industrial production by Ramey (2016) when using the Jordà method.

6 Earnings, Consumption, and Welfare

From Galí (2013) it is known that increasing wage flexibility need not increase aggregate welfare, the intuition being that rigid wages facilitate a degree of consumption smoothing that in turn stimulates activity through the aggregate demand channel. Here I consider the related question of whether the labor variety with relatively more flexibility enjoys relatively greater welfare. I begin by detailing the assumptions under which the response of aggregate consumption to shocks is equivalent to a representative agent model despite heterogeneity in wage and employment responses. I then show that, despite this equivalence, period utility is more volatile for workers with more rigid wages and therefore that the welfare in the heterogenous agent economy falls short of the output-gap equivalent representative agent economy.

Before building the model, however, it is important to clarify when the allocative wage of variety v , which I will call $W_{v,t}$, is the *UCL* for variety v and not some other function of remitted wages. Necessary assumptions are that both firms and workers have (1) accurate expectations regarding the evolution of future wages given the existing set of realized shocks and (2) access to financial markets through which they can intertemporally smooth.¹⁷ I proceed under these assumptions.

6.1 Factor Demands

I begin with the canonical Cobb-Douglass production function augmented to include two labor varieties:¹⁸

$$y_t(s) = z_t k_t(s)^\alpha (l_{1,t}(s)^\gamma l_{2,t}(s)^{(1-\gamma)})^{(1-\alpha)},$$

where k_t , $l_{v,t}$, and y_t denote capital, labor varieties $v \in \{1, 2\}$, and output at time- t , respectively. For expositional purposes, I assume that the marginal product of each labor variety is identical up to differences in their output elasticities.¹⁹ I want to consider the economy's response to aggregate demand shocks, therefore it is useful to consider these producers as producers of differentiated goods who have price setting power as in the standard New Keynesian framework thus s indexes indexes the intermediate producers differentiated good.²⁰ I assume that these producers are competitive in their input markets and take the nominal

¹⁷Note, that this does *not* preclude an implicit wage contract motivated out of differential risk aversion on the part of the worker and firm as in, for example, Rudanko (2009).

¹⁸It is without loss of generality to add additional varieties.

¹⁹Assuming heterogeneous, log-additive labor productivity – i.e. $l = en$ where l is effective units of labor produced by n physical units of labor with efficiency e – does not substantively change results.

²⁰As in the standard model, intermediate goods are aggregated into a final output good using a CES aggregation technology: $Y_t = \left(\int_0^1 y_t(s)^{\frac{\epsilon-1}{\epsilon}} \right)^{\frac{\epsilon}{\epsilon-1}}$.

input prices $W_{v,t}$, for $v \in \{1, 2\}$, and R_t as given when minimizing costs.²¹

I do not write out the price setting problem for two reasons. First, the price setting problem is entirely standard and its consequences for the path of prices and total output, as I will show, are also nearly entirely standard. Second, all results apart from the welfare consequences of wage rigidities hold in the Real Business Cycle model as well.²²

The firms choose inputs to minimize costs each period. Constant returns to scale and free flow of capital and labor across firms ensure that all firms choose the same capital to labor ratios and the same ratio of labor varieties.²³

It is then straightforward to derive the factor demands:

$$\begin{aligned} k_t(s) &= \frac{Y_t}{z_t} \left(\frac{P_t(s)}{P_t} \right)^{-\epsilon} \left(\frac{\alpha}{R_t} \right)^{(1-\alpha)} \left(\frac{W_{1,t}}{(1-\alpha)\gamma} \right)^{\gamma(1-\alpha)} \left(\frac{W_{2,t}}{(1-\alpha)(1-\gamma)} \right)^{(1-\gamma)(1-\alpha)} \\ l_{1,t}(s) &= \frac{Y_t}{z_t} \left(\frac{P_t(s)}{P_t} \right)^{-\epsilon} \left(\frac{(1-\alpha)\gamma}{W_{1,t}} \right)^{(1-(1-\alpha)\gamma)} \left(\frac{R_t}{\alpha} \right)^\alpha \left(\frac{W_{2,t}}{(1-\gamma)(1-\alpha)} \right)^{(1-\gamma)(1-\alpha)} \\ l_{2,t}(s) &= \frac{Y_t}{z_t} \left(\frac{P_t(s)}{P_t} \right)^{-\epsilon} \left(\frac{(1-\alpha)(1-\gamma)}{W_{2,t}} \right)^{(1-(1-\alpha)(1-\gamma))} \left(\frac{R_t}{\alpha} \right)^\alpha \left(\frac{W_{1,t}}{\gamma(1-\alpha)} \right)^{\gamma(1-\alpha)} \end{aligned}$$

and show that the elasticities of demand for each labor variety with respect to demand for final goods are:

$$\begin{aligned} \varepsilon_{L_1,Y} &= 1 + \Upsilon + \alpha\varepsilon_{R,Y} + (1-\alpha) [\gamma\varepsilon_{W_1,Y} + (1-\gamma)\varepsilon_{W_2,Y}] - \varepsilon_{W_1,Y} \\ \varepsilon_{L_2,Y} &= 1 + \Upsilon + \alpha\varepsilon_{R,Y} + (1-\alpha) [\gamma\varepsilon_{W_1,Y} + (1-\gamma)\varepsilon_{W_2,Y}] - \varepsilon_{W_2,Y}, \end{aligned}$$

where $\varepsilon_{L_v,Y}$, $\varepsilon_{R,Y}$, and $\varepsilon_{W_v,Y}$ for $v \in \{1, 2\}$ are the elasticities of labor demand, the rental rate, and wages with respect to demand for the final good, respectively, and Υ captures the effect of the sensitivity of the price level and relative prices to demand for the final good. Clearly, if the elasticity of wages of variety one is greater than that for labor variety two, then the elasticity of labor of variety two is larger. Differential sensitivity of wages to a shock to demand for the final good could derive either from differential wage stickiness or from differential labor supply elasticities. I remain agnostic regarding the source until Section 6.5.

²¹This is the typical assumption regulating the producer or intermediate producer in the real business cycle and New Keynesian models, respectively.

²²As an aside, this implies that heterogeneity is irrelevant in a world that is well captured by the real business cycle model.

²³To verify note: $\frac{R_t}{W_{1,t}} \frac{(1-\alpha)\gamma}{\alpha} = \frac{l_{1,t}(s)}{k_t(s)} = \frac{L_{1,t}}{K_t}$, $\frac{R_t}{W_{2,t}} \frac{(1-\alpha)(1-\gamma)}{\alpha} = \frac{l_{2,t}(s)}{k_t(s)} = \frac{L_{2,t}}{K_t u_t}$, and $\frac{W_{1,t}}{W_{2,t}} \frac{1-\gamma}{\gamma} = \frac{l_{2,t}(s)}{l_{1,t}(s)} = \frac{L_{2,t}}{L_{1,t}}$.

6.2 Earnings

Turning to earnings, it also follows that the elasticities of earnings with respect to demand for the final good are

$$\varepsilon_{E_1,Y} = \varepsilon_{E_2,Y} = 1 + \Upsilon + \alpha\varepsilon_{R,Y} + (1 - \alpha)[\gamma\varepsilon_{W_1,Y} + (1 - \gamma)\varepsilon_{W_2,Y}],$$

where $\varepsilon_{E_v,Y}$ are the elasticities of earnings with respect to demand for the final good for varieties $v \in \{1, 2\}$. The earnings elasticities are identical! This is a well-known property of the Cobb-Douglas production technology: relative expenditures (factor shares) are invariant to relative prices. Still, equal elasticities of earnings need not imply equally variable welfare, as I will show in Section 6.5.

6.3 Consumption

Consider variety-specific households that maximize the discounted value of utility flows from consumption and labor supply decisions subject to a simple budget constraint:

$$\begin{aligned} \max_{C_{v,t}, L_{v,t}, S_{v,t}} E_0 \sum_{t=0}^{\infty} \beta^t [u(C_{v,t}) - \phi v_v(L_{v,t})] \\ \text{s.t. } P_t C_{v,t} + S_{v,t+1} \leq S_{v,t}(1 + i_t) + \Pi_t + W_{v,t} L_{v,t} \end{aligned} \quad (6.1)$$

where $C_{v,t}$, $L_{v,t}$, and $S_{v,t}$ are the consumption, labor supply, and savings of the variety v household, P_t and i_t are the price level and the interest rate, and Π_t are dividends remitted from the firm. Finally, $u(\cdot)$ and $v_v(\cdot)$ are CRRA flow utilities.

This gives rise to the following Euler equation for each variety:

$$\frac{u'(C_{v,t})}{P_t} = \beta E_t \frac{u'(C_{v,t+1})(1 + i_{t+1})}{P_{t+1}} \quad (6.2)$$

In the steady state this implies that varieties will consume proportionate to their earnings whenever dividends are also proportionate to earnings and the elasticity of intertemporal substitution is identical for all varieties. For the remainder of this section, I assume this to be true.

I have shown above that, even if wages respond differentially to an increase in aggregate demand, the sensitivity of earnings is invariant across varieties. Because all varieties receive the same windfall income and because all the prices and preferences governing the Euler equation are invariant across varieties, each variety will choose to save and consume out of the windfall earnings identically. Thus, the elasticity of the consumption response is invariant

across varieties whenever dividends are distributed in proportion to earnings.

6.4 Output-gap Equivalent Representative Agent

Closing the model implies that the sum of consumption across varieties is equal to output. I can now consider a counter-factual representative agent economy in which the (elasticity of) the output gap is equivalent to the economy in which workers have heterogeneously flexible wages. Note, intermediate producers will produce equivalent output when they face equivalent costs and real price level. In the heterogeneous agent economy marginal cost is:

$$mc = \frac{1}{z} \left(\frac{R}{\alpha} \right)^\alpha \left(\frac{W_1}{(1-\alpha)\gamma} \right)^{(1-\alpha)\gamma} \left(\frac{W_2}{(1-\alpha)(1-\gamma)} \right)^{(1-\alpha)(1-\gamma)} \quad (6.3)$$

and its elasticity is:

$$\varepsilon_{mc,Y} = \alpha\varepsilon_{R,Y} + (1-\alpha) [\gamma\varepsilon_{W_1,Y} + (1-\gamma)\varepsilon_{W_2,Y}] \quad (6.4)$$

Even with heterogeneity the elasticity of consumption is invariant across varieties, the aggregate impact on the price level is identical to that which would arise in a representative agent economy with identical consumption elasticity. Further, the elasticity of the marginal cost and therefore the output gap is equivalent to the output gap in a representative worker economy with

$$\varepsilon_{W_{rep},Y} = \gamma\varepsilon_{W_1,Y} + (1-\gamma)\varepsilon_{W_2,Y} \quad (6.5)$$

where $\varepsilon_{W_{rep},Y}$ is the elasticity of wages of a representative worker. Together with the equality of the earnings elasticity, this implies

$$\varepsilon_{L_{rep},Y} = \gamma\varepsilon_{L_1,Y} + (1-\gamma)\varepsilon_{L_2,Y}. \quad (6.6)$$

So, the output-gap-equivalent representative agent has wage and labor supply elasticities that are a weighted average of the varieties with weights corresponding to the output elasticities.

6.5 Welfare

Assuming that wages are set flexibly for all varieties implies that all varieties equate their marginal rate of substitution to their marginal productivity. In this case, the welfare consequences of the output gap are identical across varieties. However, if wage flexibility holds, it must be the case that the variation across education in the cyclical and sensitivity to monetary policy shocks of wages and employment documented in the preceding sections

imply that highly educated workers' Frisch elasticity is substantially smaller than that of less-educated workers. There is limited evidence in support of this hypothesis, in part because education is often used as an instrument in identification of the representative agent's Frisch (Peterman, 2016). Alternatively, the heterogeneity documented in the preceding sections could stem from increasing rigidity of wages as education declines. Indeed, as discussed in those sections, decreasing durability of employer-employee matches as education declines implies that employers have more limited ability to smooth through transient shocks in the presence of identical near-term wage rigidities for less-educated employees.

Note that the result that all varieties experience identical earnings elasticities is invariant to the micro-foundation of differential elasticities of wages. Thus, even in an environment with nominal wage rigidities, the preceding results imply equally sensitive consumption across varieties. However, wage rigidities drive a wedge between the marginal rate of substitution and the marginal product of labor. Still, from Galí (2013) it is known that increasing wage rigidity need not decrease aggregate welfare. This derives from noting that, in a model with a representative worker variety, wage rigidity mutes the consumption response to the underlying shock, potentially resulting in a smaller fluctuation in aggregate demand. Under the assumptions made here, the intuition of Galí (2013) holds and the magnitude of the offsetting effect of wage rigidity is captured by the representative agent's elasticity of wages $\varepsilon_{W_{rep}, Y}$. In other words, variation in the elasticity of wages for the varieties improves aggregate welfare whenever the analogous variation in the output-gap equivalent $\varepsilon_{W_{rep}, Y}$ improves welfare. Despite this, however, it is straightforward to observe that period utility is 1) lower and more volatile for the variety with more rigid wages and 2) lower and more volatile when aggregated across varieties in the economy with heterogeneity as compared to the output-gap equivalent representative agent economy. These results stem from straightforward applications of Jensen's inequality.

The magnitude of the excess welfare loss in the heterogeneous agent economy relative to the representative agent economy can be assessed empirically using the framework laid out in Ball and Romer (1989) and, most closely related to this work, Galí et al. (2007). Galí et al. (2007) show that under assumptions analogous to those that govern the representative agent analogue to the model considered here welfare costs can be evaluated using data on the price and wage markups, consumption, and employment. The results above show that extending the method of Galí et al. (2007) to the heterogeneous agent economy, only requires employment to be measured at the variety level. I measure these using the CPS Monthly Surveys as in the preceding section. In addition to employment by variety, I require a measure of the output elasticity of each labor variety. In Appendix A.7, I derive a second order approximation to the welfare cost of fluctuations. In Appendix A.8, I discuss how to

Table 7: Welfare Costs of Fluctuations Relative to an Output-Gap Equivalent Representative Agent Economy (1976-2018).

	<i>Frisch Elasticity = 1</i>		<i>Frisch Elasticity = 5</i>	
	<i>EIS = 1</i>	<i>= 5</i>	<i>EIS = 1</i>	<i>= 5</i>
Heterogeneous Workers Economy				
Aggregate	0.0039	0.0590	0.0118	0.0669
	<i>1</i>	<i>1</i>	<i>1</i>	<i>1</i>
< High School	0.0100	0.0650	0.0299	0.0850
	<i>2.52</i>	<i>1.10</i>	<i>2.52</i>	<i>1.27</i>
High Sch. / Some Coll.	0.0036	0.0587	0.0109	0.0659
	<i>0.92</i>	<i>0.99</i>	<i>.92</i>	<i>0.99</i>
≥ Bachelors	0.0006	0.0557	0.0019	0.0570
	<i>0.16</i>	<i>0.94</i>	<i>0.16</i>	<i>0.85</i>
Output-Gap Equivalent Representative Worker Economy				
	0.0034	0.0584	0.0102	0.0652
	<i>0.86</i>	<i>0.98</i>	<i>0.86</i>	<i>0.98</i>

Note: Italics report the ratio to the aggregate welfare cost of fluctuations in the heterogeneous workers economy.

Source: From the USECON database I use compensation per hour (LXNFC) and real and nominal output (LXNFO and LXNFI), which refer to the nonfarm business sector; nondurable and services consumption (CNH + GSH), drawn from the respective NIPA series; and implicit price deflator (LXNFI). Unemployment and hours by educational attainment are constructed from the Current Population Survey Basic Monthly and Outgoing Rotation files, respectively. Output elasticities are recovered using the NLSY data. Author's calculations following the method of Galí et al. (2007).

recover the output elasticities from the NLSY data.

Table 7 documents the welfare costs of fluctuations relative to the costs experienced in the output-gap equivalent representative agent economy. Results depend on the modeled Frisch elasticity and intertemporal elasticity of substitution (EIS). I allow these parameters to take on values $\{1, 5\}$ and $\{1, 5\}$, respectively. As in Galí et al. (2007) welfare losses due to unemployment fluctuations are overall small.²⁴ The baseline calibration, which sets both Frisch and EIS to unity, indicates an welfare cost that exceeds the representative agent economy by more than 15 percent. Further, the welfare loss of the least educated is more than *fifteen* times larger than that of the most educated! The level of welfare losses is higher when the Frisch elasticity and EIS are larger, as noted by Galí et al. (2007), but larger

²⁴Under the baseline specification of unit elasticities, welfare costs are an order of magnitude smaller than the 0.07 benchmark of Lucas (1987). Indeed the losses presented here, which measure fluctuations in labor using the employment rate, are even smaller than Galí et al. (2007). The reason is that the employment rate varies less over the cycle than hours. However, I use employment because it can be constructed by education from 1976 onwards, whereas hours can only be constructed from 1982, as discussed in Section 2.

elasticities mute the affects of heterogeneity.

7 Conclusion

This paper documents divergence in the flexibility of wages across educational attainment. This divergence is especially evident when using a measure of wages designed to capture the allocative consequences of the long-term nature of employment relationships and the history dependence of wage remittances: the *UCL*.

I find that the wage-tenure profile of more-educated workers is differentially cyclically sensitive. In conjunction with their differentially lower separation rates, this renders allocative wages for this group more sensitive to business cycle conditions. I also document divergence in the response to monetary policy shocks across educational attainment. More educated workers' allocative wage is more sensitive while their employment is less sensitive.

I then consider these facts in an economy that features the conventional Cobb-Douglas production technology typically used in the New Keynesian modeling framework. I expand this technology to admit multiple labor varieties differentiated only by the elasticity of allocative wages with respect to an aggregate demand shock. The production technology—which is designed to replicate the Kaldor Facts—produces the strict implication that factor shares are invariant to demand and technology shocks. This holds even with price and wage rigidities and with or without constant returns to scale.

I then consider the consumption response to an aggregate demand shock when the economy consists of variety-specific households. In other words, when workers may pool income but only with workers of like labor variety. In this context it is useful to note that the factor share of each labor variety is equal to the earnings of that variety. Thus, if dividends from firm profits (if any) are remitted in proportion to earnings, consumption is equally sensitive to aggregate demand shocks regardless of variety!

In the presence of both price and wage rigidities, I show that the output-gap and price level depend only on a weighted average of the varieties wage rigidities that is a function of their factor shares. Thus, the stabilizing effect of wage rigidities, if any, described in Galí (2013) is a function only of this synthetic representative agent's wage rigidity.

Acknowledging heterogeneity, however, reveals a welfare cost of business cycle fluctuations that is more than fifteen percent larger than the cost experienced in an output-gap equivalent representative worker economy. This excess cost is borne largely by the less educated workers, whose welfare costs are more than fifteen times that of the most educated.

This leads to two conclusions. First, the welfare-maximizing Taylor-type rule should consider the unemployment gaps as well as the output gap and price level and should place

greater weight on the unemployment gap of the less-educated. Further, the skew in the weights should increase as heterogeneity increases. Second, welfare is improved when heterogeneity in wage rigidity decreases while holding constant the wage rigidity for the output-gap equivalent representative agent. These conclusions are broadly in line with recent work by [Aaronson et al. \(2019\)](#); however, accounting for cyclical sensitivity in wage-tenure profiles, as done here, reveals that cyclical sensitivities of earnings are less heterogenous across education than their estimates imply. Excess welfare losses documented here derive *only* from differential misallocation of labor over the business cycle for more and less educated workers.

However, identical consumption elasticities across varieties (which is closely related to identical earnings elasticities), stands in contrast to recent results in [Coibion et al. \(2017\)](#). Consumption equivalence stems from several assumptions of which one deviation is particularly interesting to consider: absence of financial frictions. Supposing that the allocative and remitted wage are identical in a frictionless financial market – an assumption which the empirical results here in show is counterfactual – then introducing a fixed proportion of hand-to-mouth consumers of each variety and labor variation on the extensive margin implies an increase in consumption inequality within each variety while maintaining the equivalence of the average consumption response across varieties. [Coibion et al. \(2017\)](#) do not provide a decomposition by education, so it is not straightforward to assess the veracity of this implication. However, as noted, the assumption runs contrary to the finding that the *EWV* is cyclically sensitive and generates an important portion of the cyclical sensitivity of the *UCL*. This implies that the introduction of liquidity constrained consumers attenuates the allocatively relevant wage rigidities toward the *NHW*. The welfare implications of this depend both on the consequences of hand-to-mouth consumption decisions for the price level and output-gap and on the degree of demand stabilization accrued through wage rigidities. I leave investigation of these margins for future work.

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A Appendix

A.1 Labor Market Volatility excluding the Global Financial Crisis

Table A1: Labor Market Volatility by Education (25+ years) Excluding Global Financial Crisis.

	Employment Rate ^a			Hours per Week ^b		
	Mean	Standard		Mean	Standard	
		Deviation	Volatility		Deviation	Volatility
All	6.30	1.46	1.21	38.58	0.68	0.24
< High School	12.47	2.00	1.71	34.47	0.84	0.40
High Sch. / Some Coll.	5.87	1.36	1.20	38.43	0.53	0.24
≥ Bachelors	2.68	0.56	0.46	41.30	0.44	0.28

^a *Source:* Current Population Survey Monthly Data 1976-2007 and author's calculations.

^b *Source:* Current Population Survey Outgoing Rotations 1982-2007 and author's calculations.

Note: Detrended unemployment rate reports the standard deviation of the data expressed in terms of deviations from trend. Series are detrended separately by education. Trend is recovered using the method of [Hamilton \(2018\)](#). All data are at the quarterly frequency.

A.2 Educational Upgrading On-the-job

Educational upgrading on the job potentially biases results. Table A2 shows that incidence of upgrading on the job are limited. Additionally, robustness checks (available upon request) reveal that alternative treatment of upgrading produces nearly identical results. Specifically I consider splitting the sample on contemporaneous education and on highest educational attainment achieved during the job spell. The former effectively treats an educational upgrade as a new hire, and the latter assumes that employers have ex-ante knowledge of an employees education prospects at the time of hiring.

Table A2: Educational Upgrading while Employed.

	Percent upgrading education on the job:			
	All years	1979-1988	1989-1998	1999-2008
Attain high school equivalent	1.81	3.35	0.52	0.52
Attain college degree	2.30	5.21	1.06	0.88

Source: National Longitudinal Study of Youth 1979 and author's calculations.

A.3 Hodrick-Prescott Filter vs Hamilton (2018) Filter

The Hamilton (2018) decomposition produces a smoother cycle and more volatile trend than the Hodrick-Prescott decomposition. For the data used in this paper, this is illustrated in Figure A1. A consequence is that regressions that use Hamilton (2018) filtered series as the cyclical driver produce smaller and less statistically significant coefficients.

Figure A1: Cyclical components of Hodrick-Prescott and Hamilton (2018).

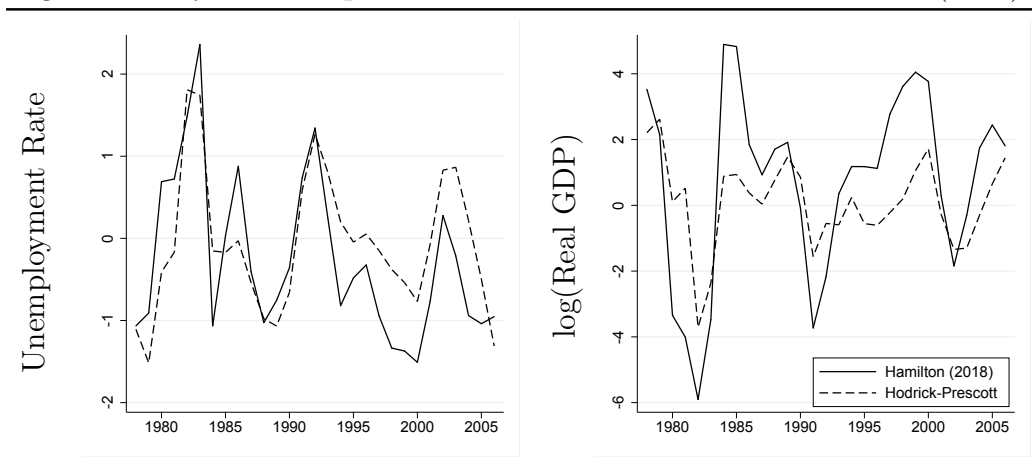


Table A3:]

Wage Cyclically by Education Group (1978-2006). [HP Filter]

Cyclical Indicator = log real GDP ^a	User Cost of Labor	New Hire's Wage	Avg. Hourly Earnings ^b
< High School	1.69 (1.14)	-0.14 (0.52)	-0.11 (0.30)
High School / Some Coll.	3.22*** (1.03)	0.68 (0.55)	0.76 (0.41)
≥ College	7.13** (1.66)	3.13*** (1.03)	1.13* (0.65)
Observations	29	29	29
Cyclical Indicator = unemployment rate ^a	User Cost of Labor	New Hire's Wage	Avg. Hourly Earnings ^b
< High School	-2.75 (1.90)	-0.55 (0.86)	-0.36 (0.51)
High School / Some Coll.	-6.35*** (1.42)	-1.66** (0.82)	-1.76** (0.59)
≥ College	-8.32* (4.91)	-5.91*** (1.29)	-2.76*** (0.83)
Observations	29	29	29

Note: All regressions control for a quadratic time trend. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Source: National Longitudinal Study of Youth 1979 and author's calculations.

^a Detrended using the HP-filter.

^b Controlling for experience, industry fixed effects, and individual fixed effects.

A.4 Coding Job Cycles and Tenure

Table A4: Reported Reasons for Separation.

Voluntary	Involuntary
<ul style="list-style-type: none"> · Quit for pregnancy, childbirth or adoption of a child · Quit to look for another job · Quit to take another job · Quit because of ill health, disability, or medical problems · Quit to spend time with or take care of children, spouse, parents, or other family members · Quit because didn't like job, boss, coworkers, pay or benefits · Quit to attend school or training · Other (SPECIFY) · Moved to another geographic area · Dissatisfied with job matching service · Sold business to another person or firm · Business temporarily inactive · Closed business or dissolved partnership 	<ul style="list-style-type: none"> · Layoff, job eliminated · Company, office or workplace closed · End of temporary or seasonal job · Discharged or fired · Government program ended · Transportation problems · Retired · No desirable assignments available · Project completed or job ended · Job assigned through a temp agency or a contract firm became permanent · Project completed or job ended · Business failed or bankruptcy · Went to jail, prison, had legal problems

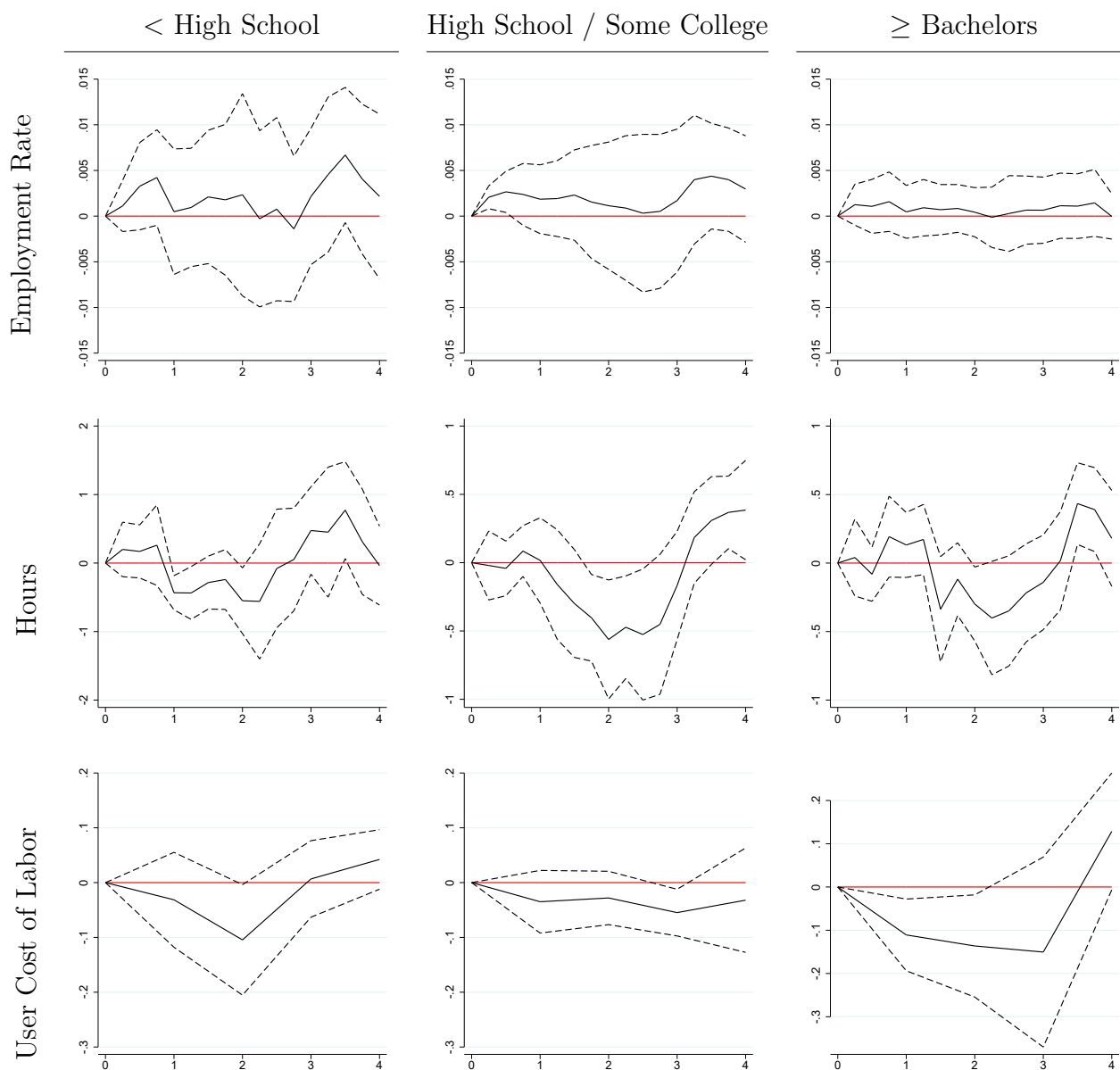
Note: Categorization follows [Basu and House \(2016\)](#) and [Hagedorn and Manovskii \(2013\)](#).

NLSY data contain weekly data on the employment situation of each individual that is collected retrospectively at the time of each (bi)yearly interview. These data record employment in up to five concurrent jobs for each week as well as the reason for termination of the employment relationship when it occurs. For each week I record information about the *primary job* defined as the job in which the worker reports the highest pay or as the job designated as primary if no pay is reported for any job. I also record the presence, if any, of secondary jobs.

An *employment cycle* is defined as a period during which a worker is continuously employed. I consider there to be a break in continuous employment if the respondent reports involuntary separation from her employer and there is no secondary job or if she reports greater than eight continuous weeks of unemployment regardless of the reason for separation. Reasons for severance are categorized into voluntary and involuntary in table A4.

Tenure is defined as the completed period during which the worker reports to have continuously worked for a given employer. Tenure is inclusive of time spent in multiple jobs such that tenure at job A begins with the first week that employer A is primary and ends with

Figure A2: Excluding Volcker Reform (1983-2007).

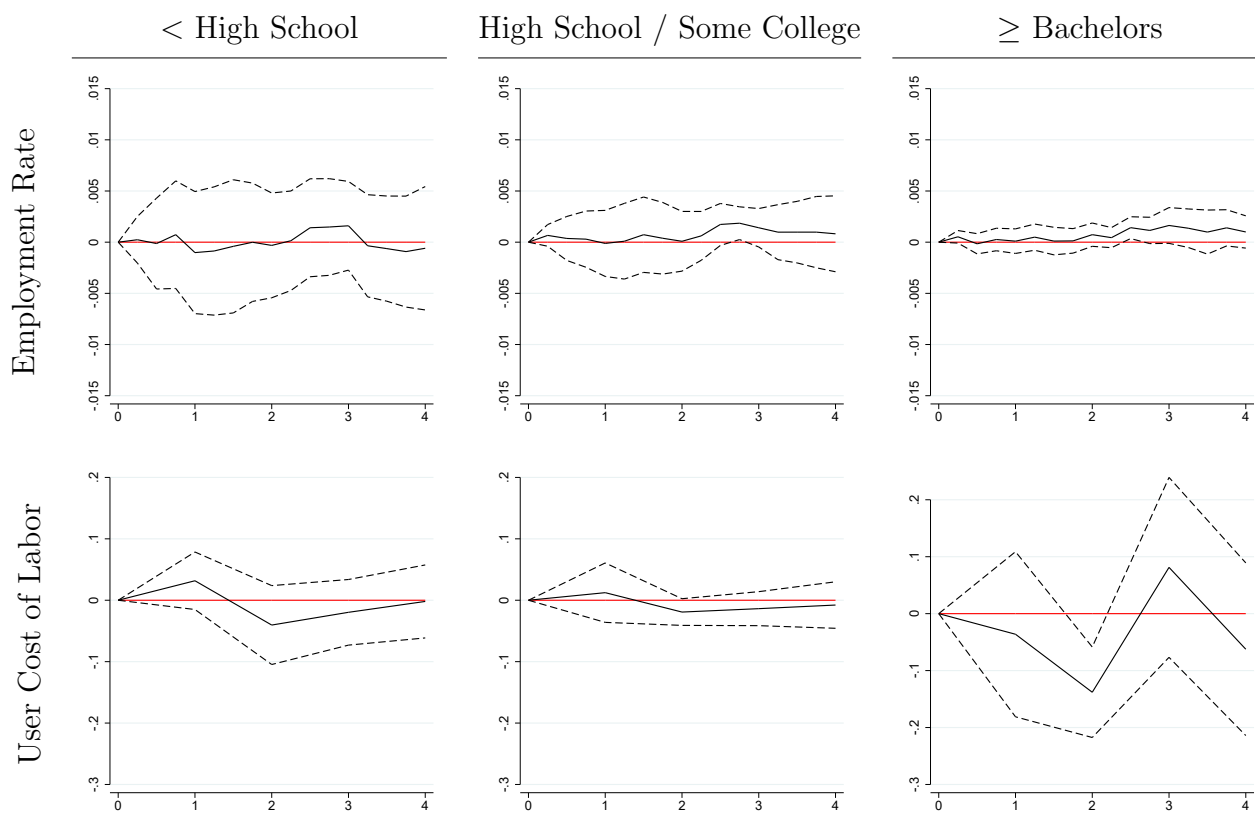


Note: 95% confidence interval.

Source: National Longitudinal Study of Youth 1979, Current Population Survey, Greenbooks as cleaned by Coibion et al. (2012), and author's calculations.

the last week that employer A is primary so long as job A was held continuously (potentially as a secondary job). Two or more discontinuous spells with the same employer result in two different tenure measures (corresponding to the completed length of each spell). Spells with employer A may be discontinuous if they are within different employment cycles or if they are interrupted by employment with another employer without employer A remaining as a

Figure A3: Estimate Shocks Excluding Pre-1979 Data.



Note: 95% confidence interval.

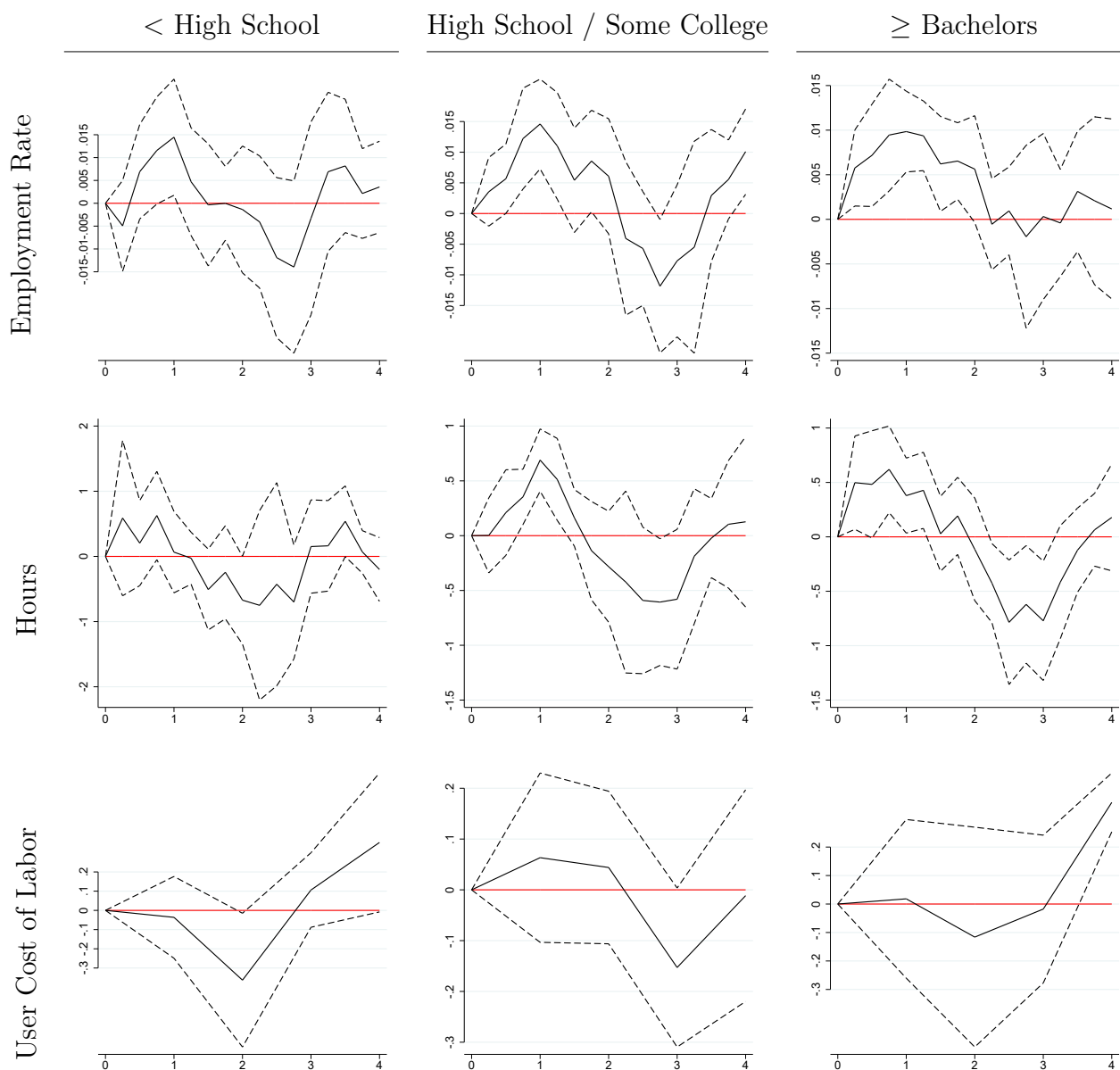
Source: National Longitudinal Study of Youth 1979, Current Population Survey, Greenbooks as cleaned by Coibion et al. (2012), and author's calculations.

secondary employer.

A.5 Alternative Identification of Monetary Policy Shocks.

I consider alternative strategies for the identification of monetary policy shocks. Figure A2 plots the impulse response of the employment rate, hours, and the user cost of labor when the Volcker Reform (179-1982) is excluded. As documented in Coibion (2012) and Ramey (2016), I find that the estimated response of employment is sensitive to inclusion of the early 1980s, although this is somewhat mitigated by hours. The second row of Figure A2 shows that hours may still be sensitive to monetary policy shocks for the least skilled even in this later period. Finally, results regarding the *UCL* are robust to exclusion of the Volcker reform. In Figure A3, I check the robustness of the identification of the underlying shocks to the inclusion of the 1969-1978 Greenbook data. Again, consistent with the literature, I find that results regarding employment are not robust. On the other hand, sensitivity of the

Figure A4: High Frequency Identification (1990-2007).



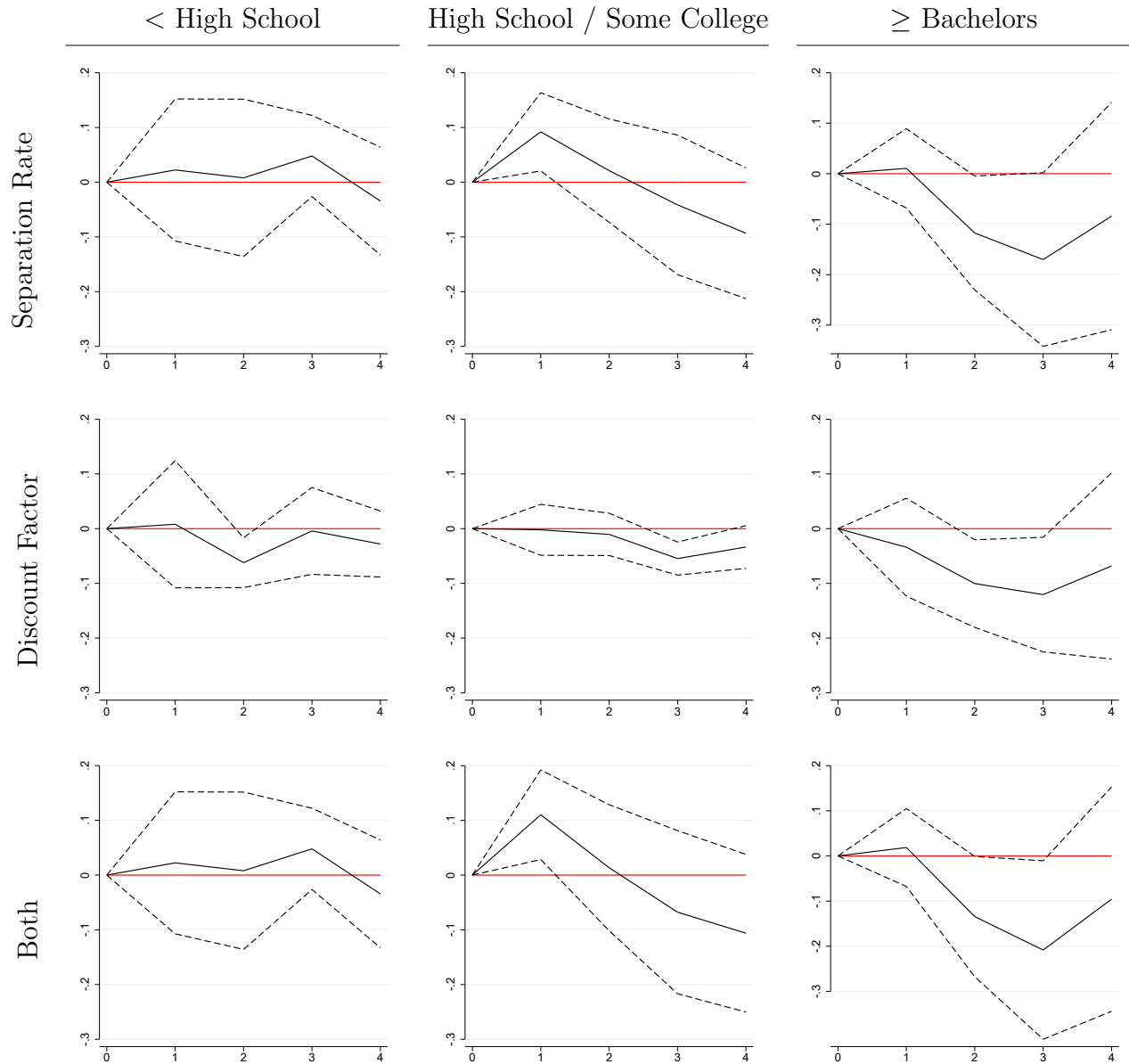
Note: 95% confidence interval.

Source: National Longitudinal Study of Youth 1979, Current Population Survey, Greenbooks as cleaned by Coibion et al. (2012), and author's calculations.

UCL to these alternatively identified shocks remains.

Finally, I consider identifying monetary policy shocks using a high frequency identification strategy as in [Gertler and Karadi \(2015\)](#). This identification strategy is possible in the period after 1990. Figure [A5](#) shows that using this identification strategy I identify “wrong signed” fluctuations in employment and hours and very large standard errors on predictions for the

Figure A5: Time-varying Separation and Discount Rates.



Note: 95% confidence interval.

Source: National Longitudinal Study of Youth 1979, Current Population Survey, Greenbooks as cleaned by Coibion et al. (2012), and author's calculations.

UCL rendering them statistically indistinguishable from zero. However, for several reasons, I place little weight on these results. As pointed out by Ramey (2016), the Gertler and Karadi (2015) shocks are (1) not mean zero, (2) autocorrelated, (3) predictable using Greenbook forecasts, and 4) provide inconsistent results when impulse responses are identified using a SVAR versus the Jordà (2005) local projection method. The finding of “wrong signed”

employment comports well with the finding of “wrong signed” industrial production by Ramey (2016) when using the Jorda method.

A.6 Time-Varying Separation and Discount Rates

Here I check the robustness of my results to allowing for time-varying separation and discount rates in the estimation of the UCL . With respect to separation rates, I allow separation rates to fluctuate over time in proportion to the fluctuations observed in the employment to unemployment transition rate recovered from the Current Population Survey and the method of Shimer (2012) while maintaining an average level consistent with that reported here for the NLSY. With respect to the discount rate, I allow the discount rate to fluctuate according to the stochastic discount factor estimated by Hall (2017). I then estimate the impulse response of these alternatively estimated UCL to a monetary policy shock estimated as in section 5.

A.7 Second Order Approximation to the Welfare Cost

Time subscripts suppressed for notational convenience.

$$Welfare\ Cost = \mathbb{E} \left[\frac{U(C, L_v) - U(\bar{C}, \bar{L}_v)}{\bar{U}_C \bar{C}} \right] \quad (A.1)$$

$$\approx \mathbb{E} \left[\frac{\bar{U}_C \bar{C} \left\{ \tilde{c} + \frac{1-\sigma}{2} \tilde{c}^2 \right\} + \bar{U}_{L_v} \bar{L}_v \left\{ \tilde{l}_v + \frac{1-\sigma}{2} \tilde{l}_v^2 \right\}}{\bar{U}_C \bar{C}} \right] \quad (A.2)$$

$$= \mathbb{E} \left[\tilde{c} + \frac{1-\sigma}{2} \tilde{c}^2 + \frac{\bar{U}_{L_v} \bar{L}_v}{\bar{U}_C \bar{C}} \left\{ \tilde{l}_v + \frac{1-\sigma}{2} \tilde{l}_v^2 \right\} \right] \quad (A.3)$$

$$\approx \left(\frac{1-\sigma}{2} \right) \mathbb{V}[\tilde{c}] - (1-\Phi) \left(\frac{1+\phi}{2} \right) \mathbb{V}[\tilde{l}_v] \quad (A.4)$$

where bars denote the value on the constant-gap path, tildes denote mean-zero log-deviations from this path, and $U(C, L_v)$ denotes the period utility enjoyed by a household of variety v consuming C and supplying labor L_v . Defining

$$(1-\Phi) = \frac{M\bar{R}S_v}{M\bar{P}L_v}, \quad (A.5)$$

the final line holds with equality if all output is consumed each period. Galí et al. (2007) argue that this is approximately true. In the present context I require, in addition, that output is consumed by varieties proportionately to their earnings, which is the implication of the Cobb-Douglas production technology and household’s intertemporal optimization

whenever Π_v is proportionate to earnings (as assumed in the main text). I follow Galí et al. (2007) and calibrate $(1 - \Phi) = \exp(-0.5)$.

A.8 Measuring the Output Elasticities of Labor Varieties

Measuring the output elasticities requires comparing the factor shares. This is complicated by the fact that allocative wages are not remitted in a given period. Thus, the factor shares must be computed from the NLSY data. I compute these as the $\{number\ employed\} * UCL$ for each educational category divided by the sum of this figure across categories. Because the NLSY is an aging cohort these shares are unstable in the early years as workers increase their educational attainment. Therefore, I compute the shares from 1986 onward. This results in shares 0.15, 0.64, and 0.21 for less than high school, high school / some college, and bachelors or more, respectively. Shares are biased toward the less educated to the extent that the less educated work fewer hours. Adjustments which account for this using hours differentials observable in the CPS data are available upon request.