Education and the Margins of Cyclical Adjustment in the Labor Market

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Abstract

I document a disparity in the cyclicality of the allocative wage–the labor costs considered when deciding to form or dissolve an employment relationship–across levels of educational attainment. Specifically, workers with a bachelor's degree or more exhibit an allocative wage that is highly pro-cyclical while high school dropouts exhibit no statistically discernible cyclical pattern. In service of this investigation, I develop new methods for inferring the cyclical sensitivity of labor costs when wages are be intertemporally smoothed or distorted and new tests to scrutinize and put to rest the possibility that variation in match quality confounds these results.

JEL Classifications:

E24:	Employment • Unemployment • Wages
J31:	Wage Differentials • Education Based • Tenure Earning
J63 & M51:	Turnover • Employee Retention
J41:	Labor Contracts • Implicit Contracts
M52:	Compensation Methods and Their Effects
E52:	Monetary Policy • Policy Effects

Keywords: User Cost of Labor, Implicit Contracts, Education and Wage Differentials, Tenure and Turnover, Wage Rigidity.

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1 Intro

1.1 Margins of adjustment by education

I document large differences in job-continuation probabilities by education.[1] Correspondingly, more educated workers have longer tenures and job-continuation probabilities are more volatile for the least educated. Differential match longevity suggests that more educated workers are more exposed to state dependence in the returns to tenure, which have been documented in a representative worker framework (Beaudry and DiNardo, 1991; Kudlyak, 2014; Basu and House, 2016). Indeed, longer expected match duration yields a strategic incentive to enter into contracts that shift remuneration inter-temporally (Thomas and Worrall, 1988; Burdett and Coles, 2003; Rudanko, 2009; Elsby, 2009). Consistent with this prior, I document stronger dependence of the wage-tenure profile on the cyclical position at the time of hiring for more educated workers. This in turn implies a greater divergence between remitted wages, both on average and for new hires, and allocative wages.[2] I quantify these two margins of adjustment using the National Longitudinal Survey of Youth (79), Panel Study of Income Dynamics, Survey of Income and Program Participation, and Current Population Survey.

1.2 Method contribution

In service of documenting these facts, I develop a new method for assessing the cyclical sensitivity of labor costs in the presence of wage contracts that are smoothed or otherwise inter-temporally distorted. The main insight is that divergence between the the UCL and the new hire's wage (NHW) stems from dependence of the return to tenure on the state at the time of hiring. I call this the expected wage wedge (EWW). As such, an estimate of the sensitivity of the UCL to a shock or a cyclical indicator can be expressed as a function of the the coefficients estimated in a Mincer-style regression augmented to admit sensitivity of the return to tenure to the state at hiring.

This method presents some advantages over that used in Kudlyak (2014); Basu and House (2016); Bils, Kudlyak and Lins (mimeo); Maruyama and Mineyama (mimeo) and oth-

[1] This work extends the analysis of Cairó and Cajner (2018) to include job-to-job mobility—adjusted for changes in the survey design as in Fujita, Moscarini and Postel-Vinay (2020). The disparity is accentuated when job-to-job flows are accounted for. (Indeed the least educated have the highest rate of separation to another job).

[2] Kudlyak (2014) demonstrates that the allocative wage in a DMP framework is the user cost of labor, defined the difference between the stream of payments in a contract entered this period and the discounted stream of payments in a contract entered next period. Basu and House (2016) demonstrate the same for the neoclassical and New Keneysian model classes.

ers. First, MORE PARSIMONIOUS REGRESSION FACILITATES DIS-AGGREGATION. INDEED, APPROACH FACILITATES dis-aggregation with RESPECT TO A CONTINUOUS CO-VARIATE. Second, the standard errors on the cyclical sensitivity of the UCL and the marginal effect of some covariate on this cyclical sensitivity follow from a straightforward implementation of the Delta method, so long as the variance-covariance matrix of the augmented Mincer regression is correct. Meanwhile, methods to obtain a correct variance-covariance matrix in the class of statistical models in which this Mincer regression sits (high dimensional fixed effects) are well studied Cameron, Gelbach and Miller (2011).[3][4]

1.3 cyclical and dynamic selection

In addition, viewing the empirical facts through the lens of cyclical sensitivity of the return to tenure facilitates interrogation of challenges to identification stemming from cyclical variation in selection into employment.

On the one hand, DISCUSS Solon, Barsky and Parker (1994) and Gertler, Huckfeldt and Trigari (2020).

On the other hand, the measuring the UCL depends upon estimating dependence of wages at later tenures on the state at hiring. But if higher quality matches are more likely to persist through to longer tenures and if opportunity to reallocate to preferred matches is pro-cyclical (for instance due to a more rapid on-the-job contact rate in the context of a job-ladder model) then estimated wage tenure-profile and its procyclciality may spuriously derive from omitting a control for match quality. Altonji and Shakotko (1987); Abraham and Farber (1987) and Topel (1991) consider the problem in a steady state and demonstrate that the omitted variable accounts for a sizeable share of the covariance between wages and tenure. Hagedorn and Manovskii (2013) extend the logic to a cyclical environment and posit that expansions afford more opportunity for workers to sort based upon match quality. This would imply that both the steady state return to tenure and the apparent cyclical sensitivity of these returns, documented by Bils (1985); Kudlyak (2014); Beaudry and DiNardo (1991), in this paper, and elsewhere, might spuriously derive from cyclical and cross-sectional variation in the omitted match quality variable. In addition, better quality matches, on account of

[3]Although it is not the focus of this paper, the collection of these innovations facilitate causal inference regarding the response of labor cost to shocks via standard applied-micro methods (difference-in-difference, instrumental variables, regression discontinuity, etc.). Specifically, it is already well understood how to obtain causal estimates of the pass-through of shocks to the wage-tenure profile. The innovations herein simply illuminate how to infer an estimate of the implied effect on the UCL at the time of the shock and the precision of this estimate.

[4]In addition, but not illustrated in this paper, the proposed methodology facilitates measuring the time of hiring at a high frequency. This facilitates estimation of the response to monetary policy shocks estimated via high-frequency identification.

their anticipated longer duration, may systematically enter into contracts that include more inter-temporal distortion, as I have documented is the case across education. However, I do not find evidence for either of these concerns in the case of match quality. This means that dynamic selection can be addressed by controlling for the omitted variable (match quality) and allowing for duration dependence in the separation probability. This verifies the solution proposed by Basu and House (2016) and is consistent with success of the solutions proposed by Bils et al. (mimeo). Meanwhile the adjustment for duration dependence increases the sensitivity of the UCL to the EWW (a result that can be shown analytically via Jensen's inequality).

NOTE: Hagedorn and Manovskii (2013) results are only obtained by the log specification (even flexible specifications can not replicate). The log specification effectively controls for the log sum of the average state during the match and tenure. I show the later to not deliver the result. Instead the correlation between the former and the state at hiring is the cause of the HM result. Basu and House (2016) control for the level, which is the more direct adjustment of Abraham and Farber (1987). WRITE ALL THIS UP IN AN APPENDIX

2 Data

Data used in this study come from three main sources: the National Longitudinal Survey of Youth, the Current Population Survey, and the Survey of Income and Program Participation. Specifics regarding the data sets are provided in this section. Following Basu and House (2016) all nominal valued variables are converted to real value using the Bureau of Economic Analysis' (BEA) Implicit Price Deflator.

Also following Basu and House (2016), I use an HP-filter to find the deviations from trend of the national unemployment rate and log real GDP as indicators of the cyclical position. These data come from the Bureau of Labor Statistics and BEA, respectively.

2.1 National Longitudinal Survey of Youth, 1979 Wave

2.2 Current Population Survey

The Current Population Survey (CPS) is a monthly rotating survey of approximately 40 thousand households upon which U.S. official employment statistics are based. Each wave quires labor force status, basic demographics, and job attributes if employed. The survey has a panel structure and supplemental topics which can be linked via an individual identifier as detailed below.

2.2.1 Labor Flows

The CPS is an eight month rotating panel separated into two four month blocks such spaced such that each respondent responds in the same calendar months in two consecutive years. In each survey respondents are asked questions used to classify them as employed, unemployed, or inactive (not in the labor force). Following Shimer (2012) and utilizing the panel identifiers supplied by Flood, King, Rodgers, Ruggles and Warren (n.d.), I link respondents over time to obtain probabilities of flow between employment, unemployment, and inactivity each month. Beginning with the 1994 survey redesign, which implemented an electronic referenced based questionnaire, respondents who are employed for two consecutive months may be asked if they hold the same job. Following Fallick and Fleischman (2004) I construct the monthly probability of job-to-job transition and following Fujita et al. (2020) I adjust these to reflect changes in the sample that are asked the same job question in later years.

2.2.2 Earnings

In the fourth and eighth interview employed respondents are asked about their weekly earnings during the reference week (the week including the 12^{th} of each month). These supplements are known as the earners study. For salaried employees, weekly earnings are queried directly and for hourly employees hourly wages and hours in the reference week are collected.[5] Weekly earnings are top-coded at \$999 prior to 1988, \$1,923 from 1989-1997, and \$2,884.61 from 1998 onward. In the main analysis, I include only individuals whose weekly earnings are below the top code.[6]

While earnings data are obtained twice for many respondents —in the fourth and eighth month in sample—potential job changes are not observed in the eight unsurveyed months interval between the fourth and fifth months in sample. As a result, tenure (discussed below) can be reliably established for at most one of the earning observations. However, the latent earning capacity of a worker and match can be proxied earnings information from the earner's study that is more remote from the tenure supplement interacted with a categorical variable describing the possibility or lack thereof that the job match is the same for both earners studies.[7]

[5] Note, this means that hours in the reference week are not collected for salaried employees. However, the core survey contains questions about usual hours in the main job. Since earnings are influenced both by pay rate and hours (even for salaried employees, who may not work 40 hours per week), I included usual hours as a control in all wage regressions.

[6] I check the robustness of point-estimates obtained via a linear regression with this restriction by replicating almost exactly in a Tobit specification. However, I prefer the censored linear specification because multi-way clustering is well studied in that context.

[7] There are four cases, 1) the tenure supplement occurs in the first four waves and there is not evidence of subsequent job changing in the basic surveys, 2) the tenure supplement occurs in the first four waves and

2.2.3 Job Tenure

While job flows can be ascertained from the core survey, the core does not contain information regarding tenures of longer durations. However, in 1983, 1987 and every other year since 1996 the CPS has fielded a supplemental module of questions regarding tenure to employed workers in either January or February.[8] In the tenure supplement, workers are asked the duration during which they have worked for their current main employer. From 1996 on, responses may report in days, weeks, months or years and are prompted to provide duration in months if they initially respond in years fewer than three. In 1983 and 1987 respondents reported in months for durations shorter than a year and years thereafter.

Tenure can be linked to the core survey and to the earnings information in the outgoing rotations via the individual identifiers (again, I rely on those verified by Flood et al. (n.d.)). Specifically, for a worker who responds to the tenure supplement in the four months leading up to and including an earners study, I infer tenure as the reported tenure advanced by the intervening time. For respondents who were surveyed during the month of an tenure supplement and were not employed at that time but were employed at the time of the earners study, I infer a tenure corresponding to the time between the most recent accession to work—as inferred from the employment data in the core survey—and the earners study. The result is 64 (16 years $\times$ 4 months per year) crosssections, in which the date at the time of the earnings information and the inferred start date of employment are observed along side demographic (age, education, sex, race, etc.) and job-attribute (industry, occupation, usual hours, hourly/salary, union status, etc.) information from the core and earners study.

2.3 Survey of Income and Program Participation, 1996-2008 Waves

The Survey of Income and Program Participation (SIPP) is an ongoing series of panels, each of moderately long duration of three to five years. The 1996, 2001, 2004, and 2008 panels share a common questionnaire and cover the years from 1996-2013. In each of these panels, surveys are conducted once every four months and information about employment, start date, and earnings information for up to two jobs in each of the intervening months is recorded.[9] For hourly workers the data record the typical pay rate. For salary earners total

there is evidence of subsequent job mobility, 3) the tenure supplement occurs in the second four waves and reported tenure is greater than a year (encompassing the first earner study), 4) the tenure supplement occurs in the second four waves and reported tenure or tenure inferred from observed job mobility when reported tenure is missing is shorter than a year.

[8]To be clear, the supplement was fielded in February 1996, 1998, and 2000 and in January in all other years. Due to the bi-yearly frequency, no respondent is eligible to have responded in more than one tenure supplement.

[9]This results in monthly data with varying degrees of recall bias. Following the literature I control for the “seam” when utilizing the SIPP panels (Ham, Li and Shore-Sheppard, 2009).

monthly earnings usual weekly hours are recorded and, following Moscarini and Postel-Vinay (2017), I infer an hourly pay rate equivalent based on the potential weeks of work in the month, reported number of weeks worked at all jobs in the month, and the worker's tenure on the job in question.[10] Data also include demographics (age, education, sex, race, etc.) and job-attributes (industry, occupation, union status, etc.).

The moderate length of the SIPP panels is an advantage of the CPS as it allows for more rigorous controls, specifically pertaining to the co-variation between tenure and match quality. The repeating, but non-overlapping, moderate length panel structure poses a unique set of (surmountable) problems in estimating the cyclical sensitivity of the UCL. Specifically, the wage-tenure profile following a realization of the macroeconomic state at time t is necessarily split between two or more SIPP panels. For example, the returns to tenure for a worker hired in 1999 will be observe from the 1996 SIPP panel for the first two years, the 2001 panel for the next three, then the 2004 panel etc.. This confounds the use of individual fixed effects or moments of the realized wage observations as proxies for match or worker quality. The problem is easy to see by considering the returns to tenure in an X start-date job: the worker's fixed effect obtained in each SIPP panel is correlated with the average tenure realized during the panel. This results in discrete declines in the estimated return to tenure given start date X at the seams between SIPP panels. I surmount this by instrumenting for tenure using the difference between realized tenure at date t and average tenure within the match for the respondent, as in Altonji and Shakotko (1987). This has the additional benefit, which should not be sold short, of addressing concerns about covariation between match quality and tenure and the business cycle.

IMPLEMENT USING IVREGHDFE woot!

2.4 Panel Study of Income Dynamics

All results hold here too. Not sure if its worth including.

2.5 Burning Glass Technologies

Address outstanding question regarding possible cyclical fluctuations in the vacancy component of the UCL.

[10]The number of weeks at each job is not queried.

3 Estimating the Sensitivity of Wages

The cyclical sensitivity of the wage is a notoriously difficult macroeconomic object to measure. Not only is there substantial qualitative and quantitative divergence between the various measures put forth, but there is also disagreement about which measure is substantively correct. Kudlyak (2014) and Basu and House (2016) argue that the appropriate measure of allocative wage to consider from the macroeconomic perspective is the UCL. This measure admits, 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.

A drawback is that the existing methodology for estimating the UCL and its sensitivity to business cycle conditions and monetary policy shocks (due to Kudlyak, 2014; Basu and House, 2016) follows a multistep procedure. This existing methodology for recovering the cyclicality of the UCL relies on (1) estimating coefficients on a very large set of indicators, which capture the return to having been hired on a specific past date given employment on a particular current date, and (2) using these coefficients to construct the time series of the UCL and then analyzing the properties of the resulting time-series. This strategy makes cross-sectional disaggregation and high-frequency measurement difficult, as both increase the already very large size of the block of indicators. In particular, investigating variation in the cyclical sensitivity of the UCL with respect to a continuous covariate is impossible, as this exercise would increase the necessary set of indicators infinitely. In addition, even in the case of a categorical covariate with few categories, inference is problematic, as relevant covariances are lost in the multistep procedure. Meanwhile, failure to address heterogeneity assumes that agents do not differ in the cyclical sensitivity of their return to tenure *or* their propensity to be observed at any given tenure horizon, potentially resulting in bias. Finally, low-frequency measurement is problematic when considering such questions as the effect of monetary policy shocks.[11]

Here, I provide a more parsimonious estimation strategy for recovering the response of the (log) UCL to a deviation in macroeconomic variable x from trend. The new strategy allows higher-frequency measurement of the cyclical position at the time of hiring and inference regarding heterogeneity in the cyclical sensitivity of the UCL—including with respect to a continuous covariate. In addition, because the procedure has fewer steps, it is easier to pinpoint the relevant covariations yielding headline results and, therefore, to interrogate

[11] Without the innovations in methodology developed in this paper, analysis of monetary policy shocks requires interpolation of data, as in Basu and House (2016).

potential sources of bias.

I begin by writing down a formulation of the wage component of the UCL augmented to admit steady state returns to tenure.

$$UCL_{e,t} = \mathbb{E}_t \sum_{j=0}^{\infty} \beta^j \left[(1-s_{e,t})^j w_{e,t,t+j} - \beta(1-s_{e,t}) \frac{w^*_{e,j+1}}{w^*_{e,j}} (1-s_{e,t+1})^j w_{e,t+1,t+j+1} \right], \quad (3.1)$$

where $w_{e,t,t+j}$ is the wage paid at date $t+j$ to a worker of type e hired on date t. $w^*_{e,j}$ is the expected wage paid to a worker of type e with tenure j who was hired in the steady state. Accounting for returns to tenure requires accounting for the change in the characteristics of the labor that accrue due to having been employed for one period.[12] If a worker with one year of tenure is not the same as a worker with zero years (and a worker with two years is not the same as a worker with one, etc), the UCL must account for these differences in worker characteristics. The $\frac{w^*_{e,j+1}}{w^*_{e,j}}$ term accounts for the changed price of the labor input due to having accrued a period of tenure.[13] Finally, $s_{e,t}$ is a type and cohort (hiring date) specific separation rate.

Now, let $\hat{x}_t$ be a log-deviation (deviation) from trend of macroeconomic indicator x. Noting that $\frac{dz}{dy} = \frac{dln(z)}{dy} z$, the (semi-)elasticity of the wage component of the user cost of labor is

$$\begin{aligned}
\mathbb{E}_t\left[\frac{dln(UCL_{e,t})}{d\hat{x}_t}\right]\Bigg|_{\hat{x}_t=0} =& \mathbb{E}_t\left[\frac{1}{UCL_{e,t}}\frac{dUCL_{e,t}}{d\hat{x}_t}\right]\Bigg|_{\hat{x}_t=0} \qquad (3.2)\\
=& \mathbb{E}_t\Bigg[\frac{1}{UCL_{e,t}}\sum_{j=0}^{\infty}\beta^j\Bigg[w_{e,t,t+j}\left((1-s_{e,t})^j\frac{dln(w_{e,t,t+j})}{d\hat{x}_t} - j(1-s_{e,t})^{j-1}\frac{ds_{e,t}}{dx_t}\right)\\
&- \beta\frac{w^*_{e,j+1}}{w^*_{e,j}}w_{e,t+1,t+j+1}\Bigg((1-s_{e,t})(1-s_{e,t+1})^j\frac{dln(w_{e,t+1,t+j+1})}{d\hat{x}_t}\\
&- (1-s_{e,t+1})^j\frac{ds_{e,t}}{dx_t} - j(1-s_{e,t})(1-s_{e,t+1})^{j-1}\frac{ds_{e,t+1}}{dx_t}\Bigg)\Bigg]\Bigg]\Bigg|_{\hat{x}_t=0}
\end{aligned}$$

Assumption 1. *There exists a steady state in which the macro indicator is expected to be on trend next period if it is on trend now:* $\mathbb{E}_t[\hat{x}_{t+1}|\hat{x}_t = 0] = 0$.

Assumption 1 implies that $\mathbb{E}_t[w_{e,t,t+j}|\hat{x}_t = 0] = \mathbb{E}_t[w_{e,t+1,t+j+1}|\hat{x}_t = 0] = w^*_{e,j}$ and $\mathbb{E}_t[s_{e,t}|\hat{x}_t = 0] =$

[12] For example, the growth rate of general human capital with respect to tenure, an explicitly back-loaded contract designed to reduce turnover, or evolving bargaining power over the course of the match.

[13] In the appendix show the version with $w^*_{e,j,t}$.

$\mathbb{E}_t[s_{e,t+1}|\hat{x}_t = 0] = s^*_e$. Plugging these into 3.1, yields

$$UCL_{e,t}\Big|_{\hat{x}_t=0} = \sum_{j=0}^{\infty} \beta^j (1-s^*_e)^j \left[w^*_{e,j} - \beta(1-s^*_e)\frac{w^*_{e,j+1}}{w^*_{e,j}} w^*_{e,j} \right] = w^*_{e,0}. \tag{3.3}$$

Assumption 1 and equation 3.2 imply

$$\begin{aligned}\frac{dln(UCL_{e,t})}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} =& \frac{1}{w^*_{e,0}} \frac{dUCL_{e,t}}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} \\ =& \frac{1}{w^*_{e,0}} \left(\sum_{j=0}^{\infty} \beta^j \mathbb{E}_t (1-s^*_e)^j \left[w^*_{e,j} \left(\frac{dln(w_{e,t,t+j})}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} - \frac{j}{1-s^*_e}\frac{ds_{e,t}}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} \right) \right.\right. \\ & \left.\left. - \beta w^*_{e,j+1} \left((1-s^*_e) \frac{dln(w_{e,t+1,t+j+1})}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} - \frac{ds_{e,t}}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} - j\frac{ds_{e,t+1}}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} \right) \right] \right)\end{aligned} \tag{3.4}$$

Assumption 2. *The effect of a deviation of the macro indicator from trend on wages at lead j and on cohort specific separation rates is independent of calendar time t.*

Assumption 2 implies $\mathbb{E}_t\left[\frac{dln(w_{e,t,t+j})}{d\hat{x}_t}\big|\hat{x}_t = 0\right] = \frac{dln(w_{e,0,j})}{d\hat{x}_0}^*$ and $\mathbb{E}_t\left[\frac{dln(w_{e,t+j+1,t+j})}{d\hat{x}_t}\big|\hat{x}_t = 0\right] = \frac{dln(w_{e,1,j+1})}{d\hat{x}_0}^*$, where $\frac{dln(w_{e,0,j})}{d\hat{x}_0}^*$ is the percentage premium over the average return to the j^{th} period of tenure for worker type e due to having been hired when the macroeconomic variable of interest was $\hat{x}$ above trend. Similarly, $\frac{dln(w_{e,1,j+1})}{d\hat{x}_0}^*$ is the percentage premium over the average return to the j^{th} period of tenure due to having been hired one period after the macroeconomic variable of interest was $\hat{x}$ above trend. Also, $\frac{ds^*_{e,0}}{d\hat{x}_0}$ and $\frac{ds^*_{e,1}}{d\hat{x}_0}$ are the separation rates for cohorts of e type workers hired at the time and one period after the time the macroeconomic variable of interest was $\hat{x}$ above trend. So, the effect of a unit deviation in $\hat{x}$ on the UCL in the neighborhood of the steady state can be written as

So, equation 3.2 simplifies to

$$\begin{aligned}\frac{dln(UCL_{e,t})}{d\hat{x}_t}\bigg|_{\hat{x}_t=0} =& \mathbb{E}_t \left(\sum_{j=0}^{\infty} \beta^j (1-s^*_e)^j \left[\frac{w^*_{e,j}}{w^*_{e,0}} \left(\frac{dln(w_{e,0,j})}{d\hat{x}_0}^* - \frac{j}{1-s^*_e}\frac{ds_{e,0}}{d\hat{x}_0}^* \right) \right.\right. \\ & \left.\left. - \beta\frac{w^*_{e,j+1}}{w^*_{e,0}} \left((1-s^*_e)\frac{dln(w_{e,1,j+1})}{d\hat{x}_0}^* - \frac{ds_{e,0}}{d\hat{x}_0}^* - j\frac{ds_{e,1}}{d\hat{x}_0}^* \right) \right] \right).\end{aligned} \tag{3.5}$$

Pause. For expositional purposes, consider for a moment the case in which there is only one type and the separation rate is stationary. In this case 3.2 simplifies to

$$\mathbb{E}_t\left[\frac{dln(UCL_t)}{d\hat{x}_t}\right]\Bigg|_{\hat{x}_t=0} = \sum_{j=0}^{\infty}\beta^j(1-s^*)^j\left[\frac{w_j^*}{w_0^*}\frac{dln(w_{0,j})}{d\hat{x}_0}^* - \beta\frac{w_{j+1}^*}{w_0^*}(1-s^*)\frac{dln(w_{1,j+1})}{d\hat{x}_0}^*\right]. \tag{3.6}$$

Also for expositional purposes, consider just the first element of the sum:

$$\underbrace{\frac{dln(w_{0,0})}{d\hat{x}}^*}_{A} - \underbrace{\frac{w_1^*}{w_0^*}}_{B}\underbrace{\beta(1-s^*)}_{C}\underbrace{\frac{dln(w_{1,0})}{d\hat{x}}^*}_{D}. \tag{3.7}$$

For now, I will assume that C is a scalar measured without error. If I were only interested in the A term, I could run a regression of log wages on $\hat{x}$ interacted with an indicator for being a new hire. The coefficient on the interaction term provides an estimate of A. If I want to know A and B, I can expand my model to contain an indicator for being a new hire and an indicator for having one year of tenure. The exponentiated difference between the coefficients on these indicators provides an estimate of B. Finally, suppose I were only interested in D, I could run a regression of log wages on the *lag* of $\hat{x}$ interacted with the indicator for being a new hire. To obtain a standard error for the entirety of 3.7, however, I need also to know the variance of *and* covariance between these point estimates.

If $\hat{x}$ and its lag were uncorrelated, I could run a regression of log wages on an indicator for being a new hire and an indicator for having one year of tenure both interacted with $\hat{x}$ and its lag and be good to go.[14] But they are correlated. Meanwhile, the UCL asks us to think about the effect of a shock today on wages in relationships formed today *and* on relationships formed next period *unconditionally*. These effects can be estimated, as well as the covariance of the resulting coefficients, by expanding the data to include an observation of wages at each tenure associated with the cyclical position at the time of hiring and another associated with the cyclical position one period before the time of hiring. Of course I will have to correct the standard errors for this non-conventional data structure and I will return to this point in a couple of paragraphs.

Now, I can obtain an estimates of each of the parts of the (semi-)elasticity from an augmented Mincer regression of the form

$$\begin{aligned} ln(wage_{i,\tau}) = \sum_{j=0}^{\infty}\Big[&\zeta_j \times \mathbb{I}_{i,\tau,j}^{tenure} + \gimel_j \times \mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=1} \times \hat{x}_{\tau-j} + \daleth_j \times \mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=0} \times \hat{x}_{\tau-j-1}\Big] \\ &+ controls_{i,\tau,j}\Xi_{i,j} + \varepsilon_{i,\tau}, \end{aligned} \tag{3.8}$$

[14] This is what I did in the original submission. Adjusting the specification to account for the multi-colinearity is the reason the magnitude of the estimated UCL have fallen.

where $\mathbb{I}_{i,\tau,j}^{tenure}$ is an indicator equal to one of i has tenure j at time τ and $\mathbb{I}_{d=1}$ and $\mathbb{I}_{d=0}$ is equal to one if the observation is the duplicate of i associated with contemporaneous shocks and lagged shocks, respectively. This yields estimates $\hat{\zeta}_j$, $\hat{\gimel}_j$, and $\hat{\daleth}_j$ of the return to the j^{th} period of tenure in steady state and the sensitivity of these returns to the cyclical position at hiring and the year before hiring, respectively. Supposing β and s^* are known, I can construct an estimate of a first order linear approximation to the (semi-)elasticity of the UCL to $\hat{x}$ in the neighborhood of $\hat{x} = 0$:

$$\frac{d\widehat{ln(UCL)}}{d\hat{x}} = \sum_{j=0}^{\infty} e^{\hat{\zeta}_j - \hat{\zeta}_0}(\beta(1-s^*))^j \hat{\gimel}_j - e^{\hat{\zeta}_{j+1} - \hat{\zeta}_0}(\beta(1-s^*))^{j+1}\hat{\daleth}_j. \tag{3.9}$$

Alternatively, I can transform the data and run an alternative augmented Mincer regression of the form

$$\begin{aligned} ln(wage_{i,\tau}) = \sum_{j=0}^{\infty} \Bigg[& \zeta_j \times \mathbb{I}_{i,\tau,j}^{tenure} + \chi_j \times \frac{\mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=1}}{(\beta(1-s^*))^j} \times \hat{x}_{\tau-j} \\ & + \psi_j \times \frac{\mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=0}}{(\beta(1-s^*))^{j+1}} \times \hat{x}_{\tau-j-1} \Bigg] \\ & + controls_{i,\tau,j}\Xi_{i,j} + \varepsilon_{i,\tau}. \end{aligned} \tag{3.10}$$

This yields estimate $\hat{\chi}_j$ of the sensitivity of the return to the j^{th} period of tenure to the cyclical position at hiring *times* the probability of surviving to the j^{th} period *times* the discounting associated with a payment j periods hence. Similarly, $\hat{\psi}_j$ is an estimate of the of the sensitivity of the return to the j^{th} period of tenure to the cyclical position one period before hiring *times* the probability of surviving to the j^{th} period *times* the probability that a worker must be replaced after one period *times* the discounting associated with a payment $j+1$ periods hence. In other words, $\hat{\chi}_j = (\beta(1-s^*))^j \hat{\gimel}_j$ and $\hat{\psi}_j = (\beta(1-s^*))^{j+1}\hat{\daleth}_j$. So,

$$\frac{d\widehat{ln(UCL)}}{d\hat{x}} = \sum_{j=0}^{\infty} e^{\hat{\zeta}_j - \hat{\zeta}_0}\hat{\chi}_j - e^{\hat{\zeta}_{j+1} - \hat{\zeta}_0}\hat{\psi}_j. \tag{3.11}$$

With homogeneity, this formulation may seem more/needlessly complicated, but it is *very* helpful in negotiating heterogeneity. In either case, standard errors are straightforward to obtain via the Delta method. However, accuracy of these requires that the standard errors in the augmented Mincer regression are correct. To this end, this regression clusters standard errors at the person (essential both due to the panel structure and because each person-year appears twice in the data) and year (Solon et al., 1994; Cameron et al., 2011).

Returning to the case with heterogeneity. Suppose e can be normalized to zero for some type. Now I can consider a baseline (semi-)elasticity of the UCL for the $e = 0$ type,

$$\mathbb{E}_t\left[\frac{dln(UCL_t)}{d\hat{x}_t}\right]\Bigg|_{\hat{x}_t=0,e=0} = \sum_{j=0}^{\infty}\beta^j(1-s_0^*)^j\left[\frac{w_{0,j}^*}{w_{0,0}^*}\frac{dln(w_{0,j})}{d\hat{x}_0}^* - \beta(1-s_0^*)\frac{w_{0,j+1}^*}{w_{0,0}^*}\frac{dln(w_{1,j+1})}{d\hat{x}_0}^*\right]. \tag{3.12}$$

and a marginal effect of e in the neighborhood of $e = 0$,

$$\begin{aligned}
&\mathbb{E}_t\left[\frac{d^2ln(UCL_t)}{d\hat{x}_t de}\right]\Bigg|_{\hat{x}_t=0,e=0} = \\
&\sum_{j=0}^{\infty}\beta^j(1-s_0^*)^j\left[\frac{w_{0,j}^*}{w_{0,0}^*}\frac{dln(w_{0,j})}{d\hat{x}_0}^*\left(\frac{dln(w_{e,j})}{de}^* - \frac{dln(w_{e,0})}{de}^*\right)\right. \\
&\qquad - \beta(1-s_0^*)\frac{w_{0,j+1}^*}{w_{0,0}^*}\frac{dln(w_{1,j+1})}{d\hat{x}_0}^*\left(\frac{dln(w_{e,j+1})}{de}^* - \frac{dln(w_{e,0})}{de}^*\right) \\
&\qquad + \frac{w_{0,j}^*}{w_{0,0}^*}\left(\frac{d^2ln(w_{e,0,j})}{d\hat{x}_0 de}^* - \frac{j}{1-s_0^*}\frac{dln(w_{e=0,j})}{d\hat{x}}^*\frac{ds_e}{de}^*\right) \\
&\qquad \left. - \beta(1-s_0^*)\frac{w_{0,j+1}^*}{w_{0,0}^*}\left(\frac{d^2ln(w_{e,1,j+1})}{d\hat{x}_0 de}^* - \frac{j+1}{1-s_0^*}\frac{dln(w_{e=0,j})}{d\hat{x}}^*\frac{ds_e}{de}^*\right)\right]
\end{aligned}$$

One can obtain an estimate of the (semi-)elasticity of the UCL to $\hat{x}$ for the $e = 0$ type as

$$\frac{dln(\widehat{UCL_{e=0}})}{d\hat{x}} = \sum_{j=0}^{\infty} e^{\hat{\zeta}_j-\hat{\zeta}_0}\hat{\chi}_j - e^{\hat{\zeta}_{j+1}-\hat{\zeta}_0}\hat{\psi}_j, \tag{3.13}$$

and the marginal effect of e in the neighborhood of $e = 0$ on this (semi-)elasticity as

$$\frac{d^2\widehat{ln(UCL)}}{d\hat{x}de} = \sum_{j=0}^{\infty}(\hat{\zeta}_{e,j}-\hat{\zeta}_{e,0})\left(e^{\hat{\zeta}_j-\hat{\zeta}_0}\hat{\chi}_j - e^{\hat{\zeta}_{j+1}-\hat{\zeta}_0}\hat{\psi}_j\right) + e^{\hat{\zeta}_j-\hat{\zeta}_0}\hat{\chi}_{e,j} - e^{\hat{\zeta}_{j+1}-\hat{\zeta}_0}\hat{\psi}_{e,j}, \tag{3.14}$$

where $\hat{\zeta}_{e,j}$, $\hat{\chi}_{e,j}$, and $\hat{\psi}_{e,j}$ are the estimated coefficients from wage regression

$$\begin{aligned} ln(wage_{i,\tau}) = \sum_{j=0}^{\infty} \Bigg[& \zeta_j \times \mathbb{I}_{i,\tau,j}^{tenure} + \chi_j \times \frac{\mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=1}}{\beta^j (1 - s_{e_i})^j} \times \hat{x}_{\tau-j} + \psi_j \times \frac{\mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=0}}{\beta^{j+1} (1 - s_{e_i})^{j+1}} \times \hat{x}_{\tau-j-1} \\ & + \zeta_{e,j} \times e_i \times \mathbb{I}_{i,\tau,j}^{tenure} + \chi_{e,j} \times e_i \times \frac{\mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=1}}{\beta^j (1 - s_{e_i})^j} \times \hat{x}_{\tau-j} \\ & + \psi_{e,j} \times e_i \times \frac{\mathbb{I}_{i,\tau,j}^{tenure} \times \mathbb{I}_{d=0}}{\beta^{j+1} (1 - s_{e_i})^{j+1}} \times \hat{x}_{\tau-j-1} \Bigg] \\ & + controls_{i,\tau,j} \Xi_{i,j} + \varepsilon_{i,\tau}, \end{aligned} \tag{3.15}$$

where e is normalized to zero for the desired type, so long as the mapping s_e is known and (approximately) linear. This setup can be extended to the case where e is categorical in the usual way. e is normalized to zero for an arbitrary category and the second two lines of equation 3.15 are repeat for each other value of e with e_i being replaced with an indicator equal to one when the value of e coincides with the category in question. In addition, e_i must be interacted with the covariates in $controls_{i,\tau,j}$ whenever failure to do so would result in spurious inference. For example, if e_i mediates both the return to tenure and the return to experience then failure to interact controls for experience with e_i will result in biased estimates of the ζ, χ, and ψ.

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