The Relative Power of Employment-to-Employment Reallocation and Unemployment Exits in Predicting Wage Growth

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In previous work (Moscarini and Postel-Vinay 2016), we investigated the hypothesis that labor demand is transmitted to wage growth primarily through Employer-to-Employer (EE) transitions. Workers often quit their jobs when they receive better offers, hence the faster these EE transitions, the higher the pace of reallocation towards high wage jobs, and the higher average wage growth. Also, the more opportunities workers have to quit, the more aggressive are their employers' wage responses, to try and retain them. We derived these predictions from the comparative statics response of steady state equilibrium to aggregate profitability shocks in the canonical Burdett and Mortensen (1998) wage posting model. Firms offer contracts and commit to deliver the resulting payments to workers, but do not renegotiate when their employees receive outside offers. As a result, job stayers never receive wage raises, unless the economy is hit by aggregate shocks which lead firms to adjust their optimal wage offers, as we showed in Moscarini and Postel-Vinay (2013). In contrast, the canonical search and matching model of Mortensen and Pissarides (1994) postulates that wages are the result of pairwise bargaining, unemployment being the worker's disagreement option. In that model, it is the exit rate from unemployment to employment (UE) that best approximates the worker's ability to extract rents from his employer. We then showed using longitudinal micro data from the Survey of Income and Program Participation (SIPP) that individual real wage growth indeed covaries much more strongly with the aggregate EE transition rate than with the aggregate UE rate.

In this paper, we continue to investigate this hypothesis. Observed wage raises can reflect either reallocation by job movers or rent extraction by job stayers following outside offers. We think of the latter type of wage raises as "inflationary", because they appear like cost shocks. Accordingly, we introduce a statistical framework to measure the covariance structure of nominal, rather than real, wage growth and labor market transitions. We ask whether a counterfactual transition rate (EE, UE etc.) that applies to each worker covaries with his wage growth, even when no individual transition occurs. This counterfactual is built by projecting actual transitions on a rich set of observable worker and job characteristics and on calendar time. In order to entertain the hypothesis of a Phillips curve, we also include in this covariance structure the unemployment rate. We interpret our results through the lens of the Sequential Auction (SA) framework introduced by Postel-Vinay and Robin (2002), which allows for renegotiation. Job stayers may obtain wage raises even in steady state, when they receive outside offers. Aggregate shocks leave the wage of a job stayer unchanged, until an outside offer actually arrives and triggers renegotiation. Then, either the worker still quits to a more productive job, or he receives a matching offer from his current employer, and stays.

I Data

We begin with a description of some basic patterns found in the data. As in Moscarini and Postel-Vinay (2016) we draw our data from the SIPP 1996-2008 panels, covering 1996-2013 with few gaps. To refine the measurement of EE transitions, we adopt the methodology and results from Fujita, Moscarini and Postel-Vinay (2017). They use job IDs, start and end dates of jobs as well weekly labor force status to construct spells of employment, non-employment, and job switches. They allow for up to one week of non-employment to classify a job change as EE. The main innovation is the imputation of

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EE transitions to individuals whose information goes missing before the end of the panel, either temporarily (gaps) or permanently (attrition). To this purpose, they run a multi-logit regression of transitions from employment to another job, to unemployment, and to non participation, on the entire sample, and use the results to impute the missing information due to gaps and attrition. In either case, the logit utilizes observable characteristics of the worker and of the last, right-censored spell, such as demographics, employer size, industry and occupation, type of worker, state of residence, as well aggregate transition rates from the monthly Current Population Survey, to capture trends and business cycles from an equally representative but different dataset. When imputing transitions over gaps, the panel nature of the SIPP allows to use information observed after the gap, such as change of household, marital status, address, that is predictive of various kinds of transitions. While the procedure cannot capture a correlation between attrition and labor market transition by unobservable worker and firm characteristics, it can at least capture its observed component. The aggregate transition rates from the matched files of the monthly CPS are themselves subject to preliminary imputation, to correct for rising rates of non-response and survey attrition, which make the commonly used series from Fallick and Fleischman (2004) increasingly inaccurate over time. Finally, they trim the first few and last months of each panel, when no respondents have a "seam" between interview periods.

We study two measures of wages: for all employed workers, nominal monthly earnings, and for hourly paid workers only (about 60% of the sample) the reported hourly pay rate. We take log changes in wages as the main variable of interest. The main reason to privilege changes over levels is that in any canonical equilibrium model of a labor market where employed workers search on the job, wages can be sticky in average level due to past commitments, while transition rates respond to job creation and are forward-looking, like wage growth. In addition, an EE transition is typically associated in the data with an unusually large and positive average change in wages, as the worker climbs the job ladder and possibly reallocates to a more productive match. We are mostly interested in separating this direct effect of EE transitions on wage growth from the indirect, unobserved rent-extraction effect due to wage renegotiation following outside offers, which puts pressure on production costs.

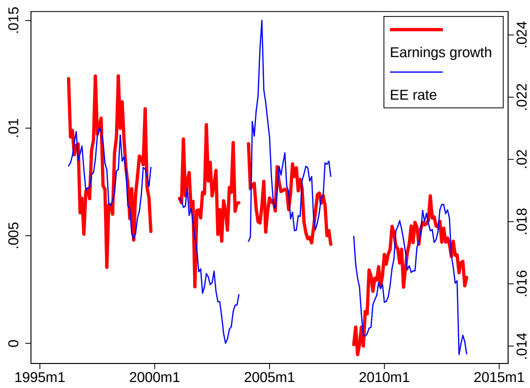
In Figure 1 we plot the log change of either monthly earnings or hourly wage rate, both in nominal terms, at monthly frequency, along with the two main transition rates of interest, EE and UE, averaged across all valid records each month, with person weights. We do it in separate graphs with EE and UE because their scales are very different. We can see common downward trends and significant pro-cyclical comovement between wage growth and job-finding rates from employment and unemployment. For earnings, the comovement is more visible with EE than with UE. For hourly wage rate, they are similar. In Figure 2 we show the same series, now passed through a linear trend. Detrending brings wage growth and EE on similar scales, hence we can plot them together with UE. The cyclical comovement between all three variables is even more visible, but EE appears to comove more strongly. To further emphasize this point, Table 1 shows results from regressions of growth in log earnings and log hourly wages on EE and UE, and clearly suggests that both measures of labor income covary more strongly with EE than with UE. The rest of this paper is dedicated to an in-depth investigation of the raw aggregate correlations reported in Table 1.

II Statistical Model

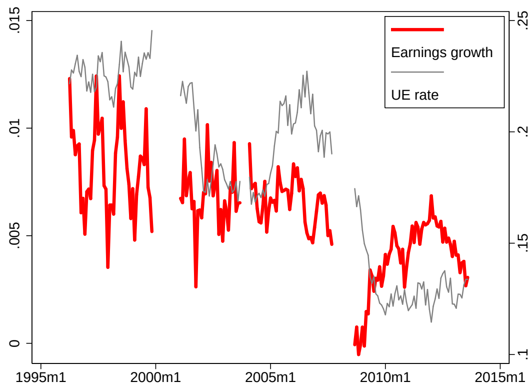
We present a flexible statistical model designed to extract the covariance over time between wage growth and various stock and flow measures of the state of the labor market. We stress that this covariance structure has no causal interpretation: our goal is to reveal time comovements of wage inflation and labor market outcomes, just as in the original Phillips curve.

From now on we refer to "wage growth" as the month-over-month change in the logarithm of either monthly nominal earnings reported by all workers, or of nominal hourly wage rate reported by hourly paid workers. We study the time-covariance structure of wage growth with various labor market transitions. To correct in a very flexible way for sample composition and eliminate cross-sectional covariance, we proceed as follows.

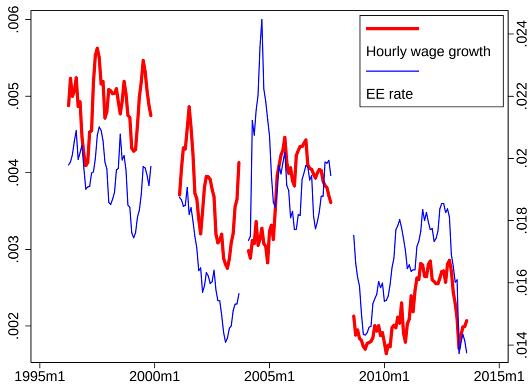
FIGURE 1. WAGE GROWTH AND TRANSITION RATES



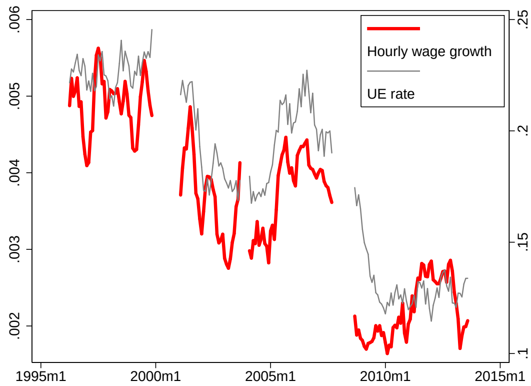
(a) Growth in nominal earnings and EE



(b) Growth in nominal earnings and UE



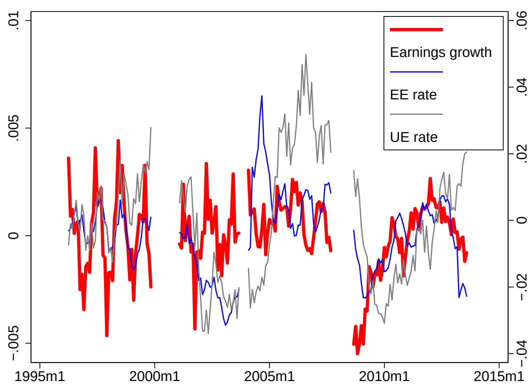
(c) Growth in nominal hourly wage rate and EE



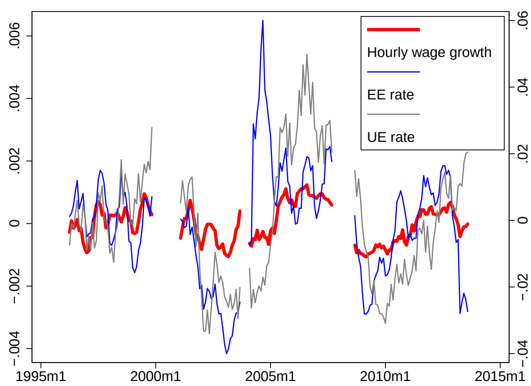
(d) Growth in nominal hourly wage rate and UE

Notes: All series are smoothed by two-quarter, two-sided MA averaging. Wage growth is on the left scale, average EE/UE transition rates on the right scale

FIGURE 2. WAGE GROWTH AND TRANSITION RATES, DETRENDED



(a) Growth in nominal earnings, EE and UE



(b) Growth in nominal hourly wage rate, EE and UE

Notes: All series are smoothed by two-quarter, two-sided MA averaging. Wage growth and EE rate are on the left scale, UE rate on the right scale.

TABLE 1—NOMINAL WAGE/EARNINGS GROWTH AND LABOR MARKET TRANSITIONS

	Dependent variable is log change in...					
	monthly nominal earnings			nominal hourly wage rate		
EE rate	0.2832		0.2834	0.1649		0.0913
	(0.0685)		(0.0741)	(0.0216)		(0.0181)
UE rate		0.0099	−0.0001		0.0215	0.0183
		(0.0067)	(0.0069)		(0.0017)	(0.0017)
# obs.			180			

Note: Monthly data, 1996m1-2013m7 (with gaps). Standard errors in parentheses. All regressions include a linear time trend. Source: SIPP and authors' calculations.

For each individual i at time t let y_{it} be either: wage growth (denoted Δw_{it}), an indicator of unemployment (U_{it}), or an indicator of labor market transition: EE_{it}, UE_{it}, EU_{it}, NE_{it}, EN_{it}. We partition the sample into cells defined by a set of K categorical individual characteristics $\left\{z_{it}^k\right\}_{k=1}^K$, where each z^k takes values in n_k classes. In what follows, we refer to a thus defined cell as a "market". Letting $\mathbb{I}$ denote the indicator function, we construct $n_1 \times \cdots \times n_K$ "market dummies"

$$m_{it}^{j_1,j_2,\cdots,j_K} = \prod_{k=1}^{K} \mathbb{I}\left(z_{it}^k = j_k\right).$$

The dummy $m_{it}^{j_1,j_2,\cdots,j_K}$ takes a value of 1 if individual i has characteristics z_{it}^k at time t that fall into classes $\{j_1, j_2, \cdots, j_K\}$, and 0 otherwise. We collate all such dummies into the vector M_{it}. We further interact those market dummies M with calendar time $s = 1,2,\cdots,T$ to construct a full set of "market-time dummies"

$$M_{it}^s = \mathbb{I}(s = t) \times M_{it}.$$

Then we proceed in two steps. In the first step we filter out observable characteristics from each variable of interest y, to eliminate composition effects. To this purpose, for each variable of interest y we run an OLS regression

$$y_{it} = X_{it}^y \beta_y + M_{it}^t \gamma_y + \varepsilon_{it}$$

with person weights, where X_{it}^y is a set of individual and job characteristics relating to the job at date $t-1$ (i.e. before the wage growth or transition occurred). Note that we partition observable worker and firm characteristics into a set X^y that enters this preliminary regression additively, without interactions, and a set M^t that enters as fully interacted dummies, including time effects.

We conclude the preliminary step by extracting the estimated market-time fixed effects for the variable of interest y:

$$\widehat{\Phi}_{it}^y = M_{it}^t \widehat{\gamma}_y$$

Note that, for any given y, each of these market-time fixed effects takes the same value for all individuals i who share the same characteristics z_{it}^k at the same time t.

In the second, main step we run an OLS regression of market-time fixed effects on each other:

$$\widehat{\Phi}_{it}^{\Delta w} = \sum_{h \in \{\mathrm{U,EE,UE,EU,EN,NE}\}} \widehat{\Phi}_{it}^h \delta_h + M_{it}\psi + \phi t + \eta_{it}$$

with person weights. This is a regression of the market-time estimated fixed effects for wage growth $\widehat{\Phi}^{\Delta w}$ on the analogous market-time fixed effects for unemployment U and transition rates EE, UE etc., a linear time trend, and a fixed effect of the market itself M_{it}, as defined by the characteristics z but not by calendar time. We think of $\widehat{\Phi}^{\Delta w}$ as a measure of latent wage growth that applies to all workers in a given market, and of the other market fixed effects $\widehat{\Phi}_{it}^h$ as counterfactual unemployment and transition rates, which estimate how likely worker i would be at time t to find another job either from employment or unemployment, or to be unemployed, based on what is currently happening to workers in the same market M_{it}^t. The estimated coefficients $\widehat{\delta}_h$ thus measure the covariance over time of wage growth with unemployment and labor market transitions within each market, after detrending and after controlling, in the preliminary step, for worker and job characteristics X.

III Empirical Implementation and Results

We define a market M as the intersection of age (5 groups), education (5), race (2) and gender (2). Adding calendar time (176 months), we have a total of $5 \times 5 \times 2 \times 2 \times 176 = 17,600$ market-time dummy variables. The linear time trend captures both trends in productivity growth and in inflation (both declined in the US over the 1996-2013 period). The additive controls X^y are, for all individuals, state of residence, for employed workers also union status, employer size (3 classes), major industry (12) and occupation (5), and a government worker dummy. This is a total of $1+2+11+4+1=19$ additional dummy covariates. We would like to "move" as many observable characteristics from the additive set X's to the fully interacted market dummies M, but we are constrained by sample size. For example, we could allow for the effect of age on the variable of interest to depend not only on education, race, gender, and calendar time, but also on state of residence. This, however, would multiply the 17,600 dummies by a further factor of 50, the number of US states. With just about 108,000 observed individual EE_{it} transitions, we would not be able to identify these covariances. We do, however, account for cross-state variation in wage growth and labor market conditions at each point in time, but in a uniform way across markets.

In the preliminary step wage growth regression we also control additively for the actual occurrence of an EE_{it} transition for that worker, to control for the direct impact of a transition on wages. In the EE transition regression we also control for the lagged level of the wage, to control for the random position of the worker in the job ladder. As expected, individual wage growth has a strong positive correlation with an actual EE_{it} transition, and in turn this transition has a strong negative correlation with lagged wage, both consistent with the logic of a job ladder model: workers who are earning a low wage, relative to their observable characteristics, are more likely to quit, and when they do quit they experience a significant wage gain. We omit the results of this first step.

The results from the second step are the main contribution of this paper. We report them in Table 2 for the change in nominal log monthly earnings, and in Table 3 for the change in the log hourly wage rate when restricting attention only to hourly paid workers.[1] In order to isolate the covariance of wage growth of job stayers only, in the last specification we also interact the $\widehat{\Phi}_{it}^{\mathrm{EE}}$ market-time fixed effect of each individual with his actual, observed EE_{it} transition.

The within-market time covariance between wage growth and EE transition rates is positive, significant, and stable across specifications. This is true even in the last column, when we control directly for observed individual EE_{it} transitions, so the baseline EE coefficient can be interpreted as pertaining only to job stayers. We also detect a "Phillips curve", a negative covariance between wage growth and unemployment rate. The estimated coefficient of the job finding rate from unemployment, UE, is consistently negative. The expected, positive sign appears instead for the NE exit rate from non participation to employment. Finally, the exit rates from employment EU and EN have negative and large estimated coefficients.

To check whether our specification of a "market" matters, we repeat the exercise only controlling for demographics and time in an additively separable manner in the first step (i.e. replacing the market-time dummy M_{it}^t by the sum of a market dummy and a time dummy in the first-step regression). We then run a regression of the estimated time effect of wage growth on estimated time effects of transitions rates and unemployment. The results are qualitatively similar.

We run these second-step regression on individuals. Those who belong in the same market-time are assigned the same, corresponding fixed effects, but are not simple duplicates of each other, because they contribute their different person weights to the regression. More generally, weights make the choice of the sample for the second step relevant. The results above refer to all individuals aged 16 to 65, whether they were employed or not. This in the spirit of interpreting these fixed effects as counterfactuals, as market averages across observationally similar workers, independently of their unobserved characteristics that correlate with employment

[1] Reported standard errors are to be taken with more than a grain of salt, because we treat the estimated fixed effects as data. We plan to bootstrap the standard errors. The very large sample size and tiny standard errors we find for now are reassuring.

TABLE 2—COVARIANCE STRUCTURE OF EARNINGS GROWTH AND LABOR MARKET TRANSITIONS

Dependent variable: log change in monthly nominal earnings							
EE rate	0.0287			0.0424	0.0383	0.0423	0.0415
	(.0006)			(.0006)	(.0006)	(.0006)	(.0006)
UE rate		-0.0004		−0.0005	−0.0011	−0.0011	−0.0011
		(.00004)		(.00004)	(.00004)	(.00004)	(.00004)
U rate			−0.0184		−0.0170	−0.0096	−0.0096
			(.0003)		(.0004)	(.0003)	(.0003)
EU rate						−0.0500	−0.0500
						(.0007)	(.0007)
NE rate						0.0257	0.0257
						(.0002)	(.0002)
EN rate						−0.0786	−0.0786
						(.0005)	(.0005)
(EE rate)*switcher							0.0312
							(.0037)
# obs.				10,784,966			

Note: Monthly data, 1996m1-2013m7 (with gaps). Standard errors in parentheses. Dependent variable and covariates are gender×ethnicity×age×education×time fixed effects estimated after controlling additively for state of residence and, for the dependent variable and Ex transition rates, also for union status, employer size (3 categories), major industry and occupation groups, and government employee status. All regressions include a linear time trend. Source: SIPP and authors' calculations.

TABLE 3—COVARIANCE STRUCTURE OF WAGE GROWTH AND LABOR MARKET TRANSITIONS

Dependent variable: monthly log change in nominal hourly wage rate							
EE rate	0.0110			0.0189	0.0177	0.0169	0.0162
	(.0002)			(.0002)	(.0002)	(.0002)	(.0002)
UE rate		−0.0015		−0.0015	−0.0019	−0.0019	−0.0019
		(.00002)		(.00002)	(.00002)	(.00002)	(.00002)
U rate			−0.0084		−0.0110	−0.0085	−0.0084
			(.0001)		(.0001)	(.0001)	(.0001)
EU rate						−0.0100	−0.0100
						(.0002)	(.0002)
NE rate						0.0045	0.0045
						(.0001)	(.0001)
EN rate						0.0034	0.0034
						(.0002)	(.0002)
(EE rate)*switcher							0.0291
							(.0012)
# obs.				10,671,642			

Note: Monthly data, 1996m1-2013m7 (with gaps), hourly paid workers only. Standard errors in parentheses. See notes to Table 2.

status. The results are qualitatively robust when we restrict attention in the second step to the workers who are employed in both the last and the current period, thus have a valid observation for wage growth.

IV Discussion

Overall, we find that the positive covariance over time between wage growth and job-to-job transition rates is supportive of a job ladder model with wage posting, with or without renegotiation. The positive comovement between wage growth of job stayers and EE transitions of observationally similar workers could be due to a common, unobserved factor, such as a shock to labor demand for that type of workers, leading firms to post more vacancies and raising opportunities for EE transitions. This common factor, however, should be captured by actual EE_{it} transitions, that we introduce in the last specification. Any residual comovement between wage growth of job stayers and the pace of EE reallocation in their labor market should only be due to rent extraction, either through pre-emptive wage raises or through matched and thus declined outside offers, that we do not directly observe. We envision this wage pressure as an important source of price pressure on firms.

The robustly small, and often negative, covariance between wage growth and UE exit rate from unemployment to employment is difficult to reconcile with the standard matching model, where wages are negotiated between workers and firms by Nash Bargaining, with unemployment being the worker outside option. Whether wage growth and UE rate respond to a common labor demand shock, or independent movements in UE affect wages through the value of the outside option, the covariance should be positive. We thus find no empirical evidence to support the view that workers, when negotiating their wages, have a credible threat to quit to unemployment, whose continuation value naturally depends on how easy it would be to then find alternative employment. Our evidence is instead consistent with a credible threat to quit, hence an ability to extract a wage raise, only once an alternative offer has arrived, or is likely to arrive soon. While UE and EE transitions strongly comove over time in response to common labor demand shocks, only the latter captures the expected pressure on wage growth.

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