How Wages Respond to the Job-Finding and Job-to-Job Transition Rates: Evidence from New Zealand Administrative Data*

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We use administrative data from New Zealand and exploit regional variations to evaluate the predictive power for wage dynamics of the job-finding and job-to-job transition rates. We find that the job-finding rate from unemployment plays a role in describing the wage dynamics of newly hired workers even after controlling for the job-to-job transition rate. The wages of new hires are much more responsive to both transition rates than the wages of job stayers. We then distinguish between the new hires transitioning from employment (job switchers) and the new hires coming from unemployment. The wages of job switchers are primarily related to the pace of job-to-job reallocation, and less significantly, to the job-finding rate. The wages of new hires from unemployment are exclusively linked to the job-finding rate and this association is stronger at the lower half of the wage distribution. Additionally, the wages of new hires from unemployment are more responsive to the job-finding rate than the wages of job stayers. The job-to-job transition rate has no impact on the wage dynamics of job stayers once the job-finding rate and the transition rate from inactivity to employment are controlled for.

JEL Codes: J31, J64.

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1. Introduction

Search-and-matching models are widely used to analyze wage dynamics. In the canonical search-and-matching model of the labor market due to Diamond, Mortensen, and Pissarides (henceforth the DMP model), the pace at which unemployed workers find jobs—the job-finding rate—is a crucial factor determining cyclical wage fluctuations. In that model, wages are set through bilateral bargaining between the employer and the employee. The bargaining power of a worker is determined by the attractiveness of their outside option, namely joining the pool of unemployed workers to look for another job. In the case of Nash bargaining, the equilibrium wage is a weighted average of a worker's productivity and her reservation wage, where the latter is directly influenced by the job-finding rate (the share of unemployed workers who transition to employment in a given period). According to the DMP model, when the job-finding rate is high, workers have more bargaining power so that, all else equal, wages increase.[1]

A key assumption in the DMP model is that unemployed workers are the only source of labor for firms to fill their vacancies. Put differently, an employed person has to first become unemployed before she can start seeking another job. This assumption ignores two groups of workers: employed people who are searching for jobs (on-the-job search) and people who are not classified as part of the labor force but who may nonetheless be willing to work and are perhaps casually looking for jobs.

Moscarini and Postel-Vinay (2016) (henceforth MPV) observe that, in the Burdett and Mortensen (1998) model (henceforth the BM model), the job-to-job transition rate is the primary driver of wages. Competition between firms for workers who are already employed drives real wages higher through two channels: a strategic effect which benefits both job stayers and job movers, and a composition effect which only benefits job movers. In the BM model, the job-finding rate plays no role in shaping wage dynamics. This prediction differs strikingly from the DMP model. Motivated by this insight, MPV analyze U.S. aggregate time-series data on the

[1]See Pissarides (2000) or Petrosky-Nadeau and Wasmer (2017) for a textbook treatment of the model.

job-finding rate and the job-switching rate.[2] MPV find that the evolution of wages over the business cycle is closely linked to the pace of job-to-job transitions and less so to variations in the job-finding rate. They interpret their findings as empirical support in favor of the BM model against the DMP model.[3]

Karahan et al. (2017) further assess the relative explanatory power of these two views of wage setting. They employ a panel data set that exploits state-level variations to measure the relative influence of the job-finding and job-to-job transition rates on cyclical wage fluctuations. They find that the wage dynamics of new hires and job stayers are both tightly linked to the pace of job-to-job transitions. Moreover, the explanatory power of the job-finding rate vanishes once they control for job-to-job flows. Their findings thus support the view that on-the-job search is the prevailing factor behind wage dynamics in the United States.[4]

Our paper is closely related to Karahan et al. (2017). We use New Zealand administrative data from the Linked Employer-Employee Data (LEED) and exploit regional variations to assess the explanatory power of the job-finding and job-to-job transition rates for wage fluctuations. We find that new hire earnings are tightly linked to the pace of job-to-job reallocation: A 1 percentage point increase in the job-to-job transition rate yields a 1.61 percent rise in the earnings of newly hired workers. However, contrary to Moscarini and Postel-Vinay (2016) and Karahan et al. (2017), we find that the job-finding rate from unemployment plays a role in describing the wage dynamics of newly hired workers even after controlling for the job-to-job transition rate. The estimated semi-elasticity of earnings to the job-finding rate is 0.30 (significant at the 5 percent level) for new hires. We also find that the wages of job stayers are much less

[2]The aggregate data were constructed by Fallick and Fleischman (2004) using the the CPS.

[3]One should note that it is the present discounted value of wages (the user cost of labor) that matters for firms' hiring decision in the DMP model. Kudlyak (2014) and Basu and House (2016) show that the user cost of labor is even more procyclical than what wage series for new hires suggest.

[4]Fallick and Fleischman (2004), Faberman and Justiniano (2015), and Mukoyama, Patterson, and Şahin (2018) provide further evidence on the importance of on-the-job search in the United States. For related evidence from New Zealand and Australia, see Karagedikli (2018) and Deutscher (2019), respectively.

responsive to the labor market conditions than the wages of newly hired workers. For job stayers, the semi-elasticities of earnings to the job-finding and job-to-job transition rates are, respectively, 0.07 and 0.22 (both significant at the 5 percent level).

We then investigate whether our baseline results remain robust when we control for the rate of transitions from inactivity to employment (the *NE* rate). Once we control for *NE* flows, the explanatory power of the job-to-job transition rate becomes insignificant for stable earnings. Thus, the strategic channel of the BM model might not be relevant in New Zealand. Job-stayer earnings then appear to be connected only to the job-finding rates from unemployment and from inactivity. This result suggests that the cyclicality of workers' attachment to the labor market, a feature overlooked in the BM and DMP models, could matter for the dynamics of stable earnings. This seems consistent with existing empirical evidence for the United States.[5] However, the LEED data do not allow us to disentangle migrants who start working quickly after their arrival in New Zealand from *NE* flows. Hence, we prefer to interpret the influence of *NE* transitions on stable earnings with a pinch of salt.

Turning to new hires, we find that controlling for the *NE* transition rate has no impact on our baseline results: the coefficients of the job-finding and job-to-job rates remain unchanged, 0.29 and 1.56, respectively (both highly significant). The coefficient of the *NE* rate is imprecisely estimated and barely significant.

Finally, to discover the source of the influence of the job-finding rate on new-hire wages, we focus on the provenance of newly hired workers. In doing so, we connect with the broader debate on the lack of wage rigidity in the data. Gertler, Huckfeldt, and Trigari (2020) argue that the pronounced procyclicality of new-hire wages observed in the data is due to workers switching jobs and does not apply to new hires coming from unemployment. The LEED data allow us to distinguish between new hires from employment (job switchers) versus new hires from unemployment. We then evaluate the predictive power of the job-finding and job-to-job transition rates for

[5] Moscarini and Postel-Vinay (2017) document a strong association between wages and the pace of *NE* flows. Elsby, Hobijn, and Şahin (2015) show that the participation margin accounts for one-third of unemployment volatility, while Armstrong and Karagedikli (2017) argue that the contribution of the participation margin could be even greater in New Zealand.

each group. We find the following results: (i) The earnings of new hires from unemployment are *exclusively* linked to the job-finding rate from unemployment, and this association is tighter at the bottom of the wage distribution. (ii) Contrary to Gertler, Huckfeldt, and Trigari (2020), the earnings of new hires from unemployment are more responsive to the labor market conditions (the job-finding rate from unemployment) than the earnings of job stayers. (iii) Consistent with a job-ladder mechanism (and with the composition effect in the BM model), the dominant predictor of job-switchers' earnings is the job-to-job transition rate, but the job-finding rate from unemployment retains some influence in the lower half of the distribution (the *NE* rate plays no role). (iv) The semi-elasticity of job-switchers' earnings to job-to-job flows is largest (equal to 3.43) at the lowest decile of the earnings distribution, highlighting the essential role of the bottom rung of the job ladder. This finding is consistent with recent evidence by Haltiwanger et al. (2018) for the United States.

2. Econometric Strategy

We first introduce the administrative data. Then we outline the empirical specifications used in our study.

2.1 Data

Our data come from a single source: the administrative Linked Employer and Employee Data (LEED) from the Inland Revenue Department (IRD).[6] LEED covers the entire population who paid some form of pay-as-you-earn (PAYE) income tax in New Zealand. We compute the number of job switchers (people switching jobs within two consecutive quarters), job stayers (continuing in the same job), unemployed people entering employment (UE), and non-participants transitioning directly into employment (NE).[7] We also calculate average nominal earnings for each of these groups.[8]

[6] The full data disclaimer is available on the last page of this paper.

[7] More specifically, LEED only enables us to classify as unemployed any person receiving an income support benefit, such as unemployment benefits.

[8] We use data on total earnings as in Karahan et al. (2017). We are unable to derive hourly earnings, as data on hours worked were not collected in LEED until April 2020.

Our quarterly observations use the middle month as the reference period, and are compared to the middle month of the previous quarter. In the New Zealand context, under the standard assumption that within-month job transitions are negligible, this approach should reduce the extent to which our quarterly data are affected by the time-aggregation bias.[9] In addition, we use the highest income employer as a reference to reduce the number of spurious job transitions identified for those with multiple employers within a month.

We then construct the time series of nominal earnings, job-finding rates, and job-to-job transition rates for each of the 16 regions in New Zealand.[10] Our data cover the 2001:Q1 to 2018:Q2 period. Figure 1 plots the job-to-job transition rate for each region. Before the global financial crisis, most regions had fluid labor markets with robust rates of job-to-job transition. Every region experienced a large decline in job-to-job flows around 2008–09, and several of them had not recovered to their pre-crisis levels by the end of our sample period.

2.2 Empirical Specifications

Our objective is similar to Karahan et al. (2017) where the empirical framework is not intended to establish a definitive causal relationship. Rather we want to distinguish between competing theories of labor market flows based on their respective prediction in terms of the association between wages, the job-finding rate, and the job-to-job transition rate.

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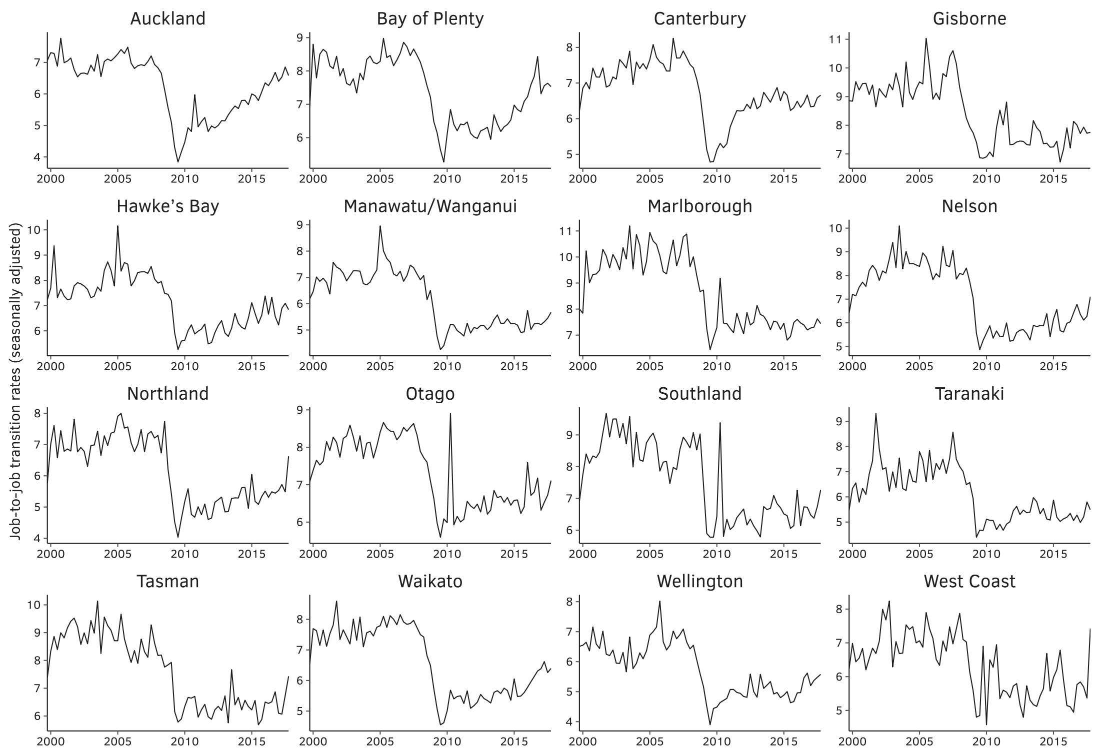
[9]In terms of our quarterly data, we may be incorrectly assigning a job-to-job transition to someone who transitions to unemployment in the first or last month of the quarter while moving between employers in the observed middle months. However, there are stand-down periods in the New Zealand transfer system which result in a delay of about three weeks before new unemployment beneficiaries receive money (in place until 2020), which lower the likelihood of these transitions being miscoded. Moreover, employees need to give their current employers at least four weeks' notice before leaving their jobs. Hence the spurious job transition events outlined in Karahan et al. (2017) seem less likely to apply to our quarterly data.

[10]Karahan et al. (2017) also use nominal earnings. There are no regional deflator data available for New Zealand.

Figure 1. Regional Job-to-Job Transition Rates in New Zealand (seasonally adjusted)



Source: Statistics New Zealand, author estimates.

We use the microdata from New Zealand to empirically test the theoretical predictions of MPV about the relative explanatory power of the job-finding and job-to-job transition rates for wage growth. In line with Karahan et al. (2017), we estimate the following two fixed-effect specifications:

$$\log W_{it} = \alpha_i + \alpha_t + \beta_i t + \alpha_u \Lambda^u_{it} + \epsilon_{it} \tag{1}$$

$$\log W_{it} = \alpha_i + \alpha_t + \beta_i t + \alpha_u \Lambda^u_{it} + \boldsymbol{\alpha_e \Lambda^e_{it}} + \epsilon_{it}, \tag{2}$$

where W_{it} denotes nominal earnings in region i in calendar quarter t. Λ^u_{it} is the transition probability from unemployment to employment in region i in quarter t. We compute it as the number of benefit recipients in region i in t–1 who enter employment in the same region within the next quarter, divided by the total number of benefit recipients in region i in t–1.[11] Λ^e_{it} is the regional job-to-job transition probability. We compute it as the share of employed people who transition from one employer in quarter t–1 to another in quarter t, with no observed intervening spell of non-employment.[12] The parameter α_i captures regional fixed effects, while the term $\beta_i t$ allows for a region-specific time trend. The time fixed effect α_t controls for variation in aggregate inflation and productivity as well as other aggregate cyclical factors. We check the validity of the fixed effects by using Hausmann's specification test. The test rejects the null hypothesis that the regional fixed effects are uncorrelated with the regressors, which confirms that our fixed-effect specification is appropriate.

Estimating specifications (1) and (2) replicates the approach of Karahan et al. (2017) for New Zealand.[13] We then include the

[11]Not all unemployed are registered to receive an unemployment benefit. Therefore, our measure of the job-finding rate is likely to exclude the very short-term unemployed, which are the most likely to get a job. This may bias our estimates of the relative predictive power of the job-finding rate downward. Below, we will explore this issue by controlling for *NE* transitions.

[12]We ensure that a person received continuous earnings throughout two consecutive quarters. We then check whether that person received income from the same employer or not.

[13]Similarly to Karahan et al. (2017), a caveat of our approach to test the predictions of the BM model is that Λ^e_{it} and Λ^u_{it} measure the realized transition rates rather than the arrival rates of job offers to employed and unemployed workers respectively.

regional transition probability from non-participation to employment, Λ_{it}^{n}, to obtain the following augmented specification:

$$\log W_{it} = \alpha_i + \alpha_t + \beta_i t + \alpha_u \Lambda_{it}^{u} + \alpha_e \Lambda_{it}^{e} + \boldsymbol{\alpha_n \Lambda_{it}^{n}} + \epsilon_{it}. \quad (3)$$

Adding the *NE* transition rate should help us to capture the very short-term unemployed (who may not register for unemployment insurance and would therefore be counted as *NE* instead of *UE* flows) and to appraise the influence of the participation margin (Elsby, Hobijn, and Şahin 2015, Armstrong and Karagedikli 2017).[14]

3. Results

We start with the replication of Karahan et al. (2017) for New Zealand. We then discuss the findings from the augmented specification that aims to account for the participation margin. Finally, we present results for the earnings distribution of job stayers, job switchers, and new hires from unemployment.

3.1 Replication of Karahan et al. (2017) for New Zealand

MPV point out that, in the BM model, wages and job-to-job transitions interact through two mechanisms: a compositional effect and a strategic rent-extraction channel. The former mechanism follows from job switchers climbing up the job ladder: workers transit between jobs when they receive (and accept) a higher wage offer. The latter effect reflects the competition between firms to retain their workforce when workers have more outside options: employees may extract a wage increase by generating an outside offer and asking their current employer to match it. In that case, the worker does not switch jobs but still gets a pay rise. Both channels favor the job-to-job transition rate over the job-finding rate from unemployment in terms of explaining the dynamics of new-hire earnings. In addition, the strategic effect also implies a positive relationship between job-to-job transitions and job-stayer earnings.

[14]It is worth noting that our LEED-based measure of *NE* flows may include recent migrants to New Zealand, which we cannot identify separately. If migrants are self-selected and there are differences in average skill sets between natives and migrants, the interpretation of the coefficient $\boldsymbol{\alpha_n}$ might not be straightforward.

Table 1. Main Results: Replication of Karahan et al. (2017)

	Stable Earners		New Hires		All	
	(1)	**(2)**	**(1)**	**(2)**	**(1)**	**(2)**
Λ^u	0.077**	0.067**	0.373**	0.300**	0.252**	0.204**
	(0.030)	(0.025)	(0.140)	(0.122)	(0.088)	(0.076)
Λ^e		0.222**		1.611***		1.054***
		(0.099)		(0.304)		(0.180)

Note: Number of observations: 1,168. Robust standard errors (clustered at regional level) in parentheses. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

The New Zealand micro-data allow us to distinguish between job stayers and new hires. Haefke, Sonntag, and van Rens (2013) show that the wages of newly hired workers are much more volatile and procyclical than the wages of job stayers. We follow Karahan et al. (2017) and estimate Equations (1) and (2) for these two groups. The results are reported in Table 1.[15] The first two columns of Table 1 report the results for stable earnings. When we only include the job-finding rate, the semi-elasticity of stable earnings to the job-finding rate is about 0.08 and significant at the 5 percent level. A 1 percentage point increase in the job-finding rate is accompanied by a 0.08 percent increase in earnings for job stayers. Controlling for the job-to-job transition rate has little impact on the explanatory power of the job-finding rate which, in contrast to Karahan et al. (2017), remains significant at the 5 percent level. The association between stable earnings and the job-to-job transition probability is stronger, with an estimated semi-elasticity equal to 0.22 and significant at the 5 percent level, providing some support for the strategic rent-extraction channel.[16]

The next two columns present the results for new-hire earnings. When included on its own, the job-finding rate is a significant

[15]We use the terms "job stayers" and "stable earners" interchangeably to denote people who remain employed in the same job.

[16]We will see in the next section that this support vanishes when we control for transitions from inactivity to employment.

predictor of earnings growth for new hires with an estimated semi-elasticity equal to 0.37. In stark contrast to Moscarini and Postel-Vinay (2016) and Karahan et al. (2017), when we control for the job-to-job transition rate, we find that the explanatory power of the job-finding rate does not evaporate: it declines slightly to 0.30 but remains significant at the 5 percent level. We find a tight link between new-hire earnings and the job-to-job transitions, with a highly significant semi-elasticity equal to 1.61. Our results confirm that the wages of newly hired workers are much more responsive to labor market conditions than the wages of job stayers, as pointed out by Haefke, Sonntag, and van Rens (2013).

The last two columns show the results for all earnings, with no distinction between new hires and stable earners. On its own, the job-finding rate is significant. Once we include the job-to-job transition rate, the explanatory power of the job-finding rate diminishes slightly, from 0.25 to 0.20, but remains significant. The coefficient estimate of the job-to-job transition rate is about five times larger than the one of the job-finding rate. A 1 percentage point increase in the job-to-job transition rate yields roughly a 1 percent rise in average earnings. To sum things up, compared to Karahan et al. (2017), the key difference is that the job-finding rate from unemployment plays a role in describing wage dynamics of newly hired workers even after controlling for the job-to-job transition rate.

3.2 Taking the Participation Margin into Account

As discussed above, the LEED data only allow us to classify as unemployed any person receiving an income support benefit, such as unemployment insurance benefit. Some of the short-term unemployed may not register for the unemployment benefit, especially the individuals who are more likely to find a job quickly (because they search harder or are more employable). Our procedure will mistakenly count these transitions as *NE* instead of *UE* flows and our measure of the job-finding rate will miss these individuals. We think this could work against the job-finding rate, and so might artificially bias the explanatory power of the job-to-job rate upward.

Controlling for the *NE* transition rate may help account for the unregistered short-term unemployed and, more generally, may shed light on the influence of the participation margin. Moscarini and

Table 2. Controlling for *NE* Transitions

	Stable Earners		**New Hires**		**All**	
	(2)	**(3)**	**(2)**	**(3)**	**(2)**	**(3)**
Λ^u	0.067**	0.060**	0.300**	0.292**	0.204**	0.197**
	(0.025)	(0.022)	(0.122)	(0.121)	(0.076)	(0.075)
Λ^e	0.222**	0.176	1.611***	1.559***	1.054***	1.004***
	(0.099)	(0.102)	(0.304)	(0.298)	(0.180)	(0.177)
Λ^n		1.441***		1.633*		1.575***
		(0.302)		(0.753)		(0.506)

Note: Number of observations: 1,168. Robust standard errors (clustered at regional level) in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Postel-Vinay (2017) find evidence of a tight link between wages and *NE* transitions in the United States. Elsby, Hobijn, and Şahin (2015) document that flows in and out of the labor force are responsible for up to a third of the cyclical volatility of the unemployment rate. Armstrong and Karagedikli (2017) argue that the significance of the participation margin might be even larger in New Zealand.

To explore these issues, we estimate specification (3), which adds the regional inactivity-to-employment transition rates. To assess the robustness of our baseline findings to controlling for *NE* transitions, Table 2 compares the results from specifications (2) and (3).

For stable earners, we find that controlling for *NE* transitions alters the estimation results in two ways. First, the influence of the job-finding rate remains stable and significant, while the predictive power of the job-to-job rate becomes insignificant. This casts doubt on the significance of the strategic rent-extraction effect in the New Zealand context. The stable earners group consists of employed people who do not engage in search activity and of employed people who engaged in on-the-job search but decided to stay in the same job. The rent-extraction channel only applies to the latter group, and it might be hard to detect evidence of it when the strategically searching job stayers are scarce compared to the non-searching job stayers. Second, the semi-elasticity of stable earnings to *NE* flows is large (equal to 1.44) and highly significant. Note, however, that the LEED data do not enable us to distinguish between new migrants and non-participants. Hence, we should interpret this coefficient cautiously. Wages of job stayers are manifestly related to the job-finding rates from inactivity and unemployment, pointing towards a link between stable earnings and workers' participation decisions.

For new-hire earnings, controlling for *NE* transitions has no effect on the baseline findings: the coefficients of the job-finding and job-to-job rates remain unchanged, at 0.29 and 1.56, respectively (both highly significant). The coefficient of the *NE* rate is imprecisely estimated and barely significant.

Interestingly, the relationship between *NE* transitions and earnings is much stronger for job stayers than for new hires. This is consistent with the view that our measure of *NE* flows partly reflect inflows of migrants who start working immediately. Coleman and Karagedikli (2018) and Howard (2020) provide evidence indicating that positive net migration flows boost aggregate demand and put broad-based upward pressures on prices throughout the economy, including the labor market, thereby lifting the average wage.

3.3 Looking at the Earnings Distribution of Stable Earners, Job Switchers, and New Hires from Unemployment

The LEED data allow us to go beyond the usual dichotomy between stable earners and new hires. To uncover the origin of the influence of the job-finding rate on new-hire wages, we follow Gertler, Huckfeldt, and Trigari (2020) and distinguish between new hires coming from unemployment (*UE* new hires) versus new hires switching from one job to another.[17] Furthermore, to understand which part of the earnings distribution contributes most to the explanatory power of the job-finding rate, we estimate specification (3) for each group at each decile of that group's earnings distribution. Table 3 shows the results for stable earners, while Table 4 focuses on the two categories of new hires.

Looking at Table 3, we see a positive relationship between stable earnings and the job-finding rate at all deciles of the earnings

[17] For completeness, we have also considered new hires transitioning from inactivity. However, none of the transition rates appeared to be linked with the earnings of that group. These results are available upon request.

Table 3. Regressions by Deciles for Stable Earnings

Decile	Job Stayers		
	Λ^u	Λ^e	Λ^n
10th	0.131*	0.722	1.154*
20th	0.081**	0.295*	1.169***
30th	0.056**	0.222*	1.334***
40th	0.041*	0.176	1.356***
50th	0.045*	0.112	1.246***
60th	0.061**	0.094	1.301***
70th	0.061**	0.103	1.299***
80th	0.059***	0.069	1.393***
90th	0.061***	0.081	1.451***

Note: Number of observations: 1,168. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

Table 4. Regressions by Deciles for New-Hire Earnings

Decile	*UE* New Hires			Job Switchers		
	Λ^u	Λ^e	Λ^n	Λ^u	Λ^e	Λ^n
10th	1.796***	1.299	−3.248	0.674**	3.430***	3.429
20th	2.040***	1.772	−2.596	0.486***	2.748***	1.558
30th	1.871***	1.165	−2.505	0.329***	2.439***	1.173
40th	1.480***	1.152	−1.749	0.249***	2.458***	1.202
50th	1.166***	1.197	−0.706	0.218***	2.496***	1.164
60th	0.931**	1.011	−0.357	0.161**	2.511***	1.455
70th	0.696**	0.918	0.334	0.096	2.495***	1.370
80th	0.404	1.031*	0.411	0.056	2.293***	0.741
90th	0.021	1.003*	0.371	0.007	1.661***	1.025

Note: Number of observations: 1,168. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

distribution. The link is quantitatively small but consistently significant. The predictive power of the job-to-job rate is insignificant at most deciles, casting doubt on the strategic effect of the BM model in the New Zealand context. Stable earnings are tightly linked with the *NE* rate. This is in line with Moscarini and Postel-Vinay (2017), who find a close relationship between wages and *NE* transitions in the United States, as well as with Armstrong and Karagedikli (2017) who argue that *NE* flows are particularly large in New Zealand. However, as discussed above, LEED does not enable us to disentangle migrants from non-participants, and we cannot rule out the conjecture that migration flows are a prime determinant of cyclical variations in stable earnings in New Zealand. A deeper analysis of the roles of migrants and non-participants in New Zealand wage dynamics is an interesting avenue for future research.

Turning to Table 4, we see that the earnings of newly hired workers transitioning from unemployment are *exclusively* linked to the job-finding rate. This association is stronger at the bottom of the distribution: the semi-elasticity is equal to 1.9 and highly significant at the three lower deciles. It then declines progressively as we move to higher deciles. Contrary to Gertler, Huckfeldt, and Trigari (2020), the earnings of new hires from unemployment are more responsive to the labor market conditions (the job-finding rate from unemployment) than the earnings of job stayers. Interestingly, the influence of the job-finding rate is stronger on new-hire wages at lower deciles of the wage distribution, in line with the findings of Katz and Krueger (1999) that the wage Phillips curve is more vivid for low-wage workers.

The earnings of job switchers are primarily connected with the job-to-job transition rate. This association is tight and highly significant (at 1 percent) throughout the entire earnings distribution. The semi-elasticity of job-switcher earnings to the job-to-job rate is 3.43 at the lowest decile. It then declines gradually to reach 1.66 at the top decile. Surprisingly, the job-finding rate also plays a significant role at the bottom half of the earnings distribution of job switchers. The explanatory power of the *NE* rate is insignificant at all deciles.

The results in this section include a number of insights, some of which are new, into the cyclical wage dynamics of newly hired workers. In contrast to Moscarini and Postel-Vinay (2016) and Karahan et al. (2017), our results show that the job-finding rate plays an important role in describing the cyclical wage dynamics of newly hired workers even after controlling for the job-to-job and *NE* rates. Indeed, the job-finding rate from unemployment is the only transition rate that consistently displays a significant relationship with the earnings of *all* categories of workers: job stayers, new hires from unemployment, and new hires switching jobs. Instead, once we control for the job-finding and *NE* rates, the job-to-job transition rate appears to be exclusively connected to the earnings of new hires switching jobs, suggestive of a pure composition effect. The fact that the semi-elasticity of job-switcher earnings to the job-to-job rate is largest at the lowest decile of the earnings distribution further highlights the importance of the bottom rung of the job-ladder, in line with Haltiwanger et al. (2018).

4. Conclusions

In this paper, motivated by recent findings of Moscarini and Postel-Vinay (2016) and Karahan et al. (2017), we have attempted to distinguish between two models of wage determination—one in which the job-finding probability of the unemployed plays a key role (the DMP model), and another in which it does not play any role for the benefit of the job-to-job transition rate (the BM model). To do so, we have used administrative data from New Zealand and replicated the empirical strategy of Karahan et al. (2017), exploiting pooled cross-regional variations to assess the explanatory power of the job-finding and job-to-job transition rate for the cyclical wage dynamics of stable earners and new hires.

Although some of our results lend support to the view that on-the-job search is a prominent factor of cyclical wage dynamics in New Zealand—specifically, new-hire earnings are tightly linked to the pace of job-to-job reallocation—some others differ starkly from the views of Moscarini and Postel-Vinay (2016) as well as Karahan et al. (2017). In particular, we find that the job-finding rate retains significant explanatory power for the earning dynamics of newly hired workers even when controlling for the pace of job-to-job transitions. This result poses a challenge to the BM model. When we control for *NE* flows, the explanatory power of the job-to-job transition rate disappears for stable earnings, suggesting that the strategic rent-extraction channel of the BM model is not relevant in the New Zealand context. Instead, the participation margin seems to matter for the evolution of stable earnings, an aspect neglected in the BM and DMP model.

To discover the source of the influence of the job-finding rate on the cyclical wage dynamics of new hires, we have then distinguished between the new hires coming from the unemployment pool and those transitioning from one job to another. For each category, we have evaluated the comparative explanatory power of the job-finding rate at different deciles of the earnings distribution. Earnings of new hires from unemployment are exclusively linked to the job-finding rate, and this association is especially tight at the bottom of the earnings distribution. Moreover, the earnings of *UE* new hires are more responsive to the labor market conditions (the job-finding rate from unemployment) than the earnings of job stayers, in contrast to evidence reported by Gertler, Huckfeldt, and Trigari (2020). Surprisingly, the job-finding rate also plays a role in describing the earnings dynamics of job switchers at the lower half of the earnings distribution. However, job switchers' earnings are primarily linked to the job-to-job transition rate, in agreement with the composition channel (i.e., a job-ladder mechanism) in the BM model. Furthermore, consistent with Haltiwanger et al. (2018), the bottom rung of the job ladder is the most essential one.

Our findings can inform recurrent policy debates on cyclical wage dynamics. Understanding the mechanisms of labor market dynamics and wage fluctuations is paramount to any central bank pursuing a flexible inflation-targeting strategy. In a number of countries, including New Zealand, the recent experience before the COVID-19 pandemic, marked by low unemployment rates and subdued wage growth left many policymakers perplexed.[18] Our results highlight the prominence of a composition effect, working through a job-ladder mechanism, for wage growth in New Zealand. Besides, our findings reveal that the job-finding rate is linked to the wage dynamics of both stable earners and new hires, especially those coming from the unemployment pool.

Finally, Kudlyak (2014) and Basu and House (2016) show that it is the present discounted value of wages that matters for firms' hiring decision in the DMP model. This is especially the case in the presence of implicit contracts whereby wages of newly hired workers respond more to the outside option than wages of existing workers (Beaudry and DiNardo 1991), as we find in the New Zealand data. An interesting avenue for future research would be to investigate how the user cost of labor (the effective price for new hires) responds to the outside option in New Zealand data.

[18] Recent papers, such as Jørgensen and Lansing (2022), argue for an important role of expectations in the behavior of the Phillips curve.

Appendix. Robustness Checks

In this appendix, we conduct a number of robustness checks for the "All Employees" category. Our first robustness check is motivated by the observation in Figure 1 that the regional job-to-job series may contain a structural break around 2008: job-to-job flows in all regions experienced a sharp fall in 2008 and, in many regions, they did not recover fully over the subsequent decade. In order to detect a possible break, we run the Bai-Perron break test. The test signals the presence of a structural break in 2008:Q2. Based on this information, we split the sample at that date and re-estimate specifications (1) to (3) over each sub-period. Table 5 displays the results. Several comments are in order. First, the influence of the job-finding rate appears to have declined over time: the point estimate of the semi-elasticity falls from 0.325 in the first period to 0.157 in the second period. Although small in magnitude, the explanatory power of the job-finding rate remains highly significant in both periods. Second, the effect of job-to-job flows seems to have become somewhat more pronounced over time: the point estimate of the semi-elasticity of earnings with respect to the job-to-job probability increases moderately from 0.878 pre-2008 to 1.175 post-2008. The explanatory power of the job-to-job transition rate is highly significant in both periods. Finally, the influence of the inactivity-to-employment transition rate is highly unstable across the two periods. This striking instability calls for some caution when interpreting the influence of *NE* flows on earnings dynamics in New Zealand. As pointed out in the main text, a caveat of our analysis is that we cannot distinguish between *NE* flows and immigrants to New Zealand who start working shortly after their arrival. We leave that issue for future research.

To further explore the nature of the structural break, we perform a second robustness check. Instead of splitting the sample period in 2008, we now introduce a dummy variable, *break*, that takes the value 0 before 2008:Q2 and 1 afterwards. The dummy variable enters in specification (2) and (3) as an interaction term with the job-finding rate Λ^u, the job-to-job transition rate Λ^e, and the inactivity-to-employment transition rate Λ^n. The results are reported in Table 6. For both specifications, the slope interaction terms are insignificant, suggesting that the explanatory power of each transition rate stays relatively constant across the two periods.

Table 5. Pre- and Post-Break Estimation

	Pre-2008				Post-2008		
$\boldsymbol{\Lambda^u}$	0.392*** (0.134)	0.326** (0.131)	0.325** (0.129)	$\boldsymbol{\Lambda^u}$	0.195*** (0.061)	0.168*** (0.050)	0.157*** (0.049)
$\boldsymbol{\Lambda^e}$		0.880*** (0.184)	0.878*** (0.180)	$\boldsymbol{\Lambda^e}$		1.226*** (0.303)	1.175*** (0.301)
$\boldsymbol{\Lambda^n}$			0.039 (0.731)	$\boldsymbol{\Lambda^n}$			2.691*** (0.530)

Note: Number of observations: 1,168. Robust standard errors (clustered at regional level) in parentheses. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

Table 6. Full Sample Estimates with Break Interactions

	(2)	(3)
Λ^u	0.261** (0.110)	0.247** (0.114)
Λ^e	1.032*** (0.167)	1.025*** (0.188)
Λ^n		0.907 (0.881)
Break	0.626*** (0.016)	0.600*** (0.024)
$\Lambda^u \times$ Break	0.083 (0.060)	–0.078 (0.068)
$\Lambda^e \times$ Break	–0.023 (0.200)	–0.123 (0.248)
$\Lambda^n \times$ Break		1.489 (1.713)

Note: Number of observations: 1,168. Robust standard errors (clustered at regional level) in parentheses. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

The break appears to be significant, but that comes from the intercept, not from the slope coefficients. Interestingly, the influence of Λ^n is now insignificant. Altogether, the two robustness checks convey the impression that our findings regarding the respective explanatory power of the job-finding rate and the job-to-job transition rate are reasonably robust, and that we should interpret the influence of *NE* flows on earnings with a pinch of salt.

Table 7. Controlling for Lagged Earnings

	(2)		(3)	
	(a)	(b)	(d)	(d)
Λ^u	0.204**	0.215**	0.197**	0.207**
	(0.076)	(0.082)	(0.075)	(0.081)
Λ^e	1.054***	1.037***	1.004***	0.998***
	(0.179)	(0.187)	(0.177)	(0.182)
Λ^n			1.575***	1.530***
			(0.505)	(0.486)
lw(−1)		0.105*		0.096
		(0.059)		(0.057)

Note: Number of observations: 1,168. Robust standard errors (clustered at regional level) in parentheses. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

Our last robustness check tackles a different issue. For the sake of comparability with Karahan et al. (2017), our baseline specifications did not include the lagged log wage. These regressions all featured time fixed effects which were expected to capture the persistent element in wages. We now check this conjecture by including the lagged log wage in specifications (2) and (3). Looking at Table 7, we see that the lagged log wage term turns out to be economically and statistically insignificant, and its inclusion does not noticeably affect the coefficients of the job-finding rate and job-to-job transition rate. This confirms the adequacy of the baseline specifications with time fixed effects.

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Access to the anonymized data used in this study was provided by Statistics NZ in accordance with security and confidentiality provisions of the Statistics Act 1975. Only people authorized by the Statistics Act 1975 are allowed to see data about a particular person, household, business, or organization, and the results in this paper have been confidentialized to protect these groups from identification.

Careful consideration has been given to the privacy, security, and confidentiality issues associated with using administrative and survey data in the IDI. Further detail can be found in the privacy impact assessment for the Integrated Data Infrastructure available from www.stats.govt.nz.

The results are based in part on tax data supplied by Inland Revenue to Statistics NZ under the Tax Administration Act 1994. This tax data must be used only for statistical purposes, and no individual information may be published or disclosed in any other form, or provided to Inland Revenue for administrative or regulatory purposes.

Any person who has had access to the unit-record data has certified that they have been shown, have read, and have understood section 81 of the Tax Administration Act 1994, which relates to secrecy. Any discussion of data limitations or weaknesses is in the context of using the IDI for statistical purposes, and is not related to the ability of these data to support Inland Revenue's core operational requirements.