The Impact of Minimum Wages on Labour Market Transitions

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Abstract

We investigate differences in labour market transition rates in high versus low minimum wage regimes using Canadian data spanning 1979 to 2008. The data include consistent questions on job tenure and reason for job separation for the whole period. Over the same time frame, there were over 140 minimum wage changes in Canada. We find that higher minimum wages are associated with lower hiring rates but also with lower job separation rates. Importantly, the reduced separation rates are due mainly to reductions in layoffs, occur in the first 6 months of a job, and are present for unskilled workers of all ages. Our estimates imply that a 10% increase in the minimum wage generates a 3.9% reduction in the layoff rate. We present a search and matching model that fits with these patterns and test its implications. Overall, our results imply that jobs in higher minimum wage regimes are more stable but harder to get. For older workers, these effects almost exactly offset each other, resulting in little impact on the employment rate. One might conclude from the small impact of minimum wages on the employment rate that they do not affect the labour market for older workers but our results indicate this is not true.

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1 Introduction

A voluminous literature exists on the impacts of minimum wages on the labour market. This literature can be seen as having two overlapping goals. The first is to use a purportedly exogenous shift in the price of a key factor to better understand the demand for labour and production decisions more generally. The second is to consider the usefulness of minimum wages as a policy tool. Almost the entire existing empirical literature on minimum wages examines the comparative static effect of a minimum wage change on employment levels and/or the shape of the wage distribution.[1] In this paper, we investigate the underlying question: how do labour market transition rates (quits, layoffs and hires) differ in low versus high minimum wage regimes? Answering this question provides a different set of insights on minimum wages as a policy tool and a new set of facts that sharpen our understanding of the functioning of the labour market.

Recent studies of employment impacts of minimum wages take one of two main approaches. The first is to compare employment levels or rates across jurisdictions with different minimum wages using panel data at the jurisdiction level (e.g., Baker et al. (1999) for Canada, Neumark and Wascher (2007) and the many papers cited therein for the US). The second is to use individual level panel data to examine the impact of an increase in the minimum wage from m_t at time t to m_{t+1} at time t+1. In particular, these latter papers examine the employment rate in t+1 for workers whose wage lies between m_t and m_{t+1} in period t (the group of workers most directly affected by the minimum wage increase). The minimum wage effect is identified by comparing employment changes for the directly affected workers with those for workers in other jurisdictions and at other points in the wage distribution (e.g., Currie and Fallick (1996) and Neumark et al. (2004) for the US; Yuen (2003) and Campolieti et al. (2005) for Canada). Both types of studies tend to find small (negative or positive) effects on employment. Our examination is closest in nature to the second of these two approaches since we study transition rates. However, we differ from those studies in three ways. First, we examine transition rates in periods before and after minimum wage increases; not transitions at the time of a change. Thus, returning to our example where the minimum wage increases between t and t+1, impacts measured in the second type of study includes layoffs between times t and t+1, while we compare quit, layoff and hiring rates between t-1 and t (i.e., in the low

[1] See Card and Krueger (1995) and Neumark and Wascher (2007) for comprehensive surveys of the literature.

minimum wage regime) to quit, layoff and hiring rates between t+1 and t+2 (the high minimum wage regime). In fact, to highlight our focus, we "dummy out" transitions spanning a minimum wage increase (transitions between times t and t+1 in this example). In addition, we focus on workers with under a year of job tenure. This means we are explicitly not trying to follow the set of workers directly affected at the time of a hike through their careers. Instead, we are investigating whether new hires who are hired after a minimum wage increase has occurred are treated differently from new hires in lower minimum wage regimes. We also differ from previous longitudinal studies in that they examine whether an affected worker is employed in the subsequent period, regardless of whether they remain with the same employer, while we focus on separations from a given employer. Our results could differ from those in these earlier studies to the extent that both separations and hires change with the minimum wage. Finally, we present estimated impacts on minimum wage changes on flows between not-in-the-labour-force (N) and unemployment (U) states, which potentially provide evidence on labour supply impacts. As far as we know, no previous paper has examined impacts on these flows.

The data we use for this exercise is the Canadian Labour Force Survey (LFS). The LFS is a representative, national survey whose main purpose is to generate data for official labour force statistics and is similar in nature to the US Current Population Survey (CPS). Importantly for us, the LFS contains a consistent question on job tenure asked in every month dating back to 1976. Like the US CPS, the LFS is actually conducted as a series of short, rolling panels with survey respondents being interviewed in each of six consecutive months. By linking data for an individual across months, we can construct monthly separation rates - the probability a job in existence in, say, March of a year, has ended by April - conditional on the duration of the job up to the initial month. We can also construct monthly hiring rates as the probability a non-employed individual in March has a new job in April, as well as NU and UN rates. We construct these transition rates separately by province and match movements in the rates to movements in the real minimum wage between 1979 and 2008, taking advantage of the very considerable variation in minimum wages across time and provinces in Canada over this period. We focus on male and female workers aged 15 to 59 with a high school or less education since we believe the minimum wage has less relevance for higher educated workers.

Remarkably, our estimates imply an economically substantial and statistically significant decrease in separation rates for low-skilled workers who have been employed for under a year in response to a minimum wage increase. In particular, a 10% increase in the real minimum wage is associated with approximately a 5% decline in the probability a worker separates from his or her job in the next year. In contrast, separation rates for workers with over a year of job tenure do not vary with the minimum wage. When we delineate by type of separation, we find that both quit and layoff rates are lower in high minimum wage regimes but that layoff rates decline more than quit rates and play a larger role in the overall reduction in separation rates. We also find that hiring rates are lower in high minimum wage regimes. Both the separation and hiring rate effects are present for workers of all ages, though the reduction in hiring is larger for teenagers. Together, this implies that the standard finding that minimum wage changes have little or no impact on employment rates for workers who are older than teenagers is a reflection of offsetting reductions in hiring and layoffs. We use our estimated hiring and separation rate effects in an equilibrium employment rate formula to show that this is indeed the case in our sample. Thus, minimum wage increases change the labour market equilibrium for workers of all ages to one with greater job security but also lower hiring rates. This result parallels findings in Blanchard and Portugal (2001) who show that although unemployment rates are similar between the US and Portugal in their data period, Portugal had much lower flows in both directions between employment and unemployment. They argue that this fits with stricter employment protection in Europe but our results imply that the same result could follow from higher minimum wages.

The other interesting empirical result is that high minimum wages are associated with lower NU transition rates and higher UN transition rates. This may imply that, for workers on the margin of participating in the labour force, an increase in the minimum wage is welfare decreasing: they place more weight on the lower probability of finding a job than on the higher wages and longer job tenures once they have a job.

The main goal of the paper is to establish these impacts of minimum wages on transition rates. These results are interesting, in part, because of the light they can shed on models of the labour market. In the fourth section of the paper, we ask whether some standard models of labour demand can rationalize the patterns we observe. We argue that Burdett-Mortensen type models, the canonical Mortensen-Pissarides search and bargaining model, and the latter model extended to allow firms to screen workers at different stages each matches some feature of the patterns we observe, but fails in fitting at least one significant feature. We argue, further, that matching the estimated pattern of effects on transition rates requires that minimum wage impacts are concentrated in the first few months of a new job, and that at least some firms do not have their value of a vacancy driven to zero by competition from new entrants. We then present a variant on a standard Mortensen-Pissarides model which has these features. In the model, match-specific productivity is revealed after the worker has been with the firm for a probationary period (say, 6 months). Once the match quality is revealed, the firm decides whether to layoff the worker. For some matches (call them minimum wage matches), productivity will be high enough that the firm does not want to terminate the match, but low enough that a bargained wage would be below the minimum wage. Such matches earn the firm lower profits than higher productivity ones, and the profitability of such matches is negatively related to the minimum wage. In this scenario, when the minimum wage rises, firms may be less willing to terminate existing non-minimum wage matches because their outside value from doing so (the profitability from finding a new match, which might end up being a minimum wage match) has declined. We present the model, in part, to see if it can generate further empirical implications. We show that under this model, firms should be less concerned about the impact of minimum wages on the expected profitability of future matches in high inflation periods where the real value of the minimum wage is declining. We find that this implication of the model is strongly confirmed in the data.

We are aware of two other papers that examine transition rates in a manner similar to what we present. Portugal and Cardoso (2006) use rich worker-firm data to look at separations and hires of teenagers before and after a 1987 increase in the Portuguese minimum wage. They also find a decline in separation rates offset by a decline in hiring. They identify their effects by comparing teenagers with older workers and argue that their results may fit with a Burdett-Mortensen type model of the labour market. Our results support their findings using a stronger identification strategy stemming from over 140 minimum wage increases, with identification coming from within-province over-time variation. In work carried out coincident to ours, Dube et al. (2012) use pairs of counties across state borders in the US between 2000 and 2008 to examine minimum wage impacts on transition rates as well as on earnings and employment levels.[2] They also find significant negative

[2]Dube et al. (2012) find that controlling for county pairs rather than just jurisdiction fixed effects reduces the size of their estimated effects. Given this, our results may be upper bounds on the estimated effects, though it is worth noting that their separation and hiring effect results remain significant when county pair controls are included.

effects of the minimum wage on hiring and separation rates, particularly for teenagers and in the restaurant industry. They then use their estimates in the context of a frictional on-the-job-search model to establish a novel measure of the importance of labour market frictions, concluding that these frictions are very important. There are two key distinctions between our paper and these two. First, we delineate quits and layoffs. Both papers focus on models in which workers quitting jobs to obtain higher wages play an important role, which does not fit well with our finding that much of the reduction in separations occurs through reduced layoffs.[3] We acknowledge that separations are usually mutually agreed upon by workers and firms in equilibrium models, muddying the distinctions between quits and layoffs. But it is difficult to believe that workers accepting outside offers from other firms would be labeled as layoffs in any data. Nonetheless, the mechanism proposed by Dube et al. (2012) may well provide an explanation for the part of the separation effect that does occur through quits, and their finding on the importance of frictions is clearly relevant for our analysis. Second, we observe worker education and use this to define low skilled labour markets. Because of this, we are able to show that minimum wage effects on transition rates are similar for older and younger workers. In the other papers, the low skilled labour market is defined as relating to teenagers and so they do not find this result.

Finally, there is a literature that examines minimum wage effects on turnover and wage distributions in the context of structural estimation (e.g., Flinn (2006) and Van den Berg (2003)). These papers adopt a much different empirical strategy relying on few minimum wage changes. We view our work as complementary to those papers in the sense that it indicates directions where further structural modeling may be useful.

The paper proceeds in five sections, including the introduction. In the second section, we describe our data. In the third section, we present our empirical strategy and the main results. In section four, we present a brief theoretical model to aid in understanding the empirical results and present a further specification indicated by the model. Section five contains conclusions.

[3]Dube et al. (2012) show that in their data over 70% of separations are quits and argue that models with quits should be the focus. But this turns out to be an artifact of their measuring separations in the restaurant industry and including students. Quits make up a third of all separations in the province of Ontario in 2007 in our sample of high school or less educated workers in all industries and excluding students. When we restrict attention to the restaurant industry and include students, the proportion that are quits rises to 68%.

2 Data

This section contains a brief description of the two main sources of data: provincial minimum wage data and Canadian LFS data. We also present basic patterns of the key variables of interest.

We use provincial minimum wage data that cover the 1979-2008 period. The minimum wage falls under provincial jurisdiction in Canada.[4] Having each of Canada's ten provinces set their own minimum wage thus provides for a rich source of minimum wage variation. Some provinces have, at various times, adopted lower rates for special classes of workers (e.g. students in Ontario). Yet, the evidence shows that firms do not, for the most part, take advantage of these special categories (e.g., Card and Krueger (1995), Shannon and Beach (1995)). As such, this paper focusses on the general adult minimum wage for each province. To match our other data, we focus on monthly frequencies. In particular, we use the minimum wage in force on the 15th of each month as relevant for that month since tenure information is asked in the week which includes the 15th of the month.

The key explanatory variable in our regression analysis is the real minimum wage. We construct it by deflating the (nominal) minimum wage by the CPI for the same province and month. Figures 1 through 3 show the real minimum wage patterns by province and year.[5] Importantly, the minimum wage shows considerable variation over time within each of the provinces.

The second source of data is the Canadian LFS master files. The LFS is a large Canadian household survey involving interviews with approximately 50,000 households per month. The focus of the LFS is to gather information on labour market activities of Canadians. A critical variable for this study comes from the LFS tenure question which asks, "When did . . . start working for his current employer". Based on the answer to this question, the LFS records the number of months of employment. What distinguishes the LFS from other Canadian data sets, and American data sets for that matter, is that this question (with no change in wording) has been asked every month since 1976.[6]

We restrict our LFS sample to individuals aged 15 to 59 with a high school or less education over

[4]Workers under federal jurisdiction (e.g. air transport) were the exception. Prior to 1996, there was a distinct federal minimum wage for those workers. The federal minimum wage was relevant to only a small subset of workers, and since 1996, the federal rate has adopted the general adult minimum wage of the province where the employer is usually employed.

[5]For ease of presentation, the figures only show the real minimum wage as of March 15th of each year (unlike our regression analysis where we use all months).

[6]See Brochu (2006) for a detailed discussion of the limitations of other North American data sets.

the 1979-2008 period.[7] We focus on high school or less educated individuals since this is the (broad) labour market for which the minimum wage is most relevant. We further exclude full-time students, the self-employed, and those in the military. Full-time students are not part of the study because working is not their main activity. The self-employed and those working in the military are removed because the processes that generate their job tenure spells are very different from (non-military) paid employees. Although LFS data is available as of 1976, we restrict our sample to January 1979 onwards to match our real minimum wage data. Provincial CPI data used to construct the real minimum wage variable is only available as of September 1978. It is worth noting that of our low skilled sample, pooled across all years and including all tenures, approximately 3% have wages within 10 cents of the relevant minimum wage. This increases to 9% for those with under 6 months of job tenure. Selecting all real minimum wage changes of at least 50 cents, 6.7% of workers in the relevant province/month had wages greater than or equal to the old minimum wage and less than or equal to the new minimum wage in the month before the change. Thus, minimum wages have a significant "bite" and could be expected to affect a non-trivial fraction of workers and firms.

Our main focus in the empirical work is on transition rates in and out of employment. To construct those rates we take advantage of the rotating panel design of the LFS. Individuals remain in the sample for six consecutive months, and every month one-sixth of the panel is replaced. As such, one can link consecutive months of the LFS thereby creating two-month mini panels.[8] With mini panels, the estimation of transition rates is straightforward. The March 2008 layoff rate for Ontario, for example, is estimated using only the March 2008 mini panel (i.e. the linked March-April 2008 data); it is simply the (weighted) proportion of period 1 Ontario workers who were identified as laid-off in period 2 of the panel.[9]

The first variable we investigate is the separation rate, defined as the probability a person on a job in month t is no longer working for that employer in month t+1, i.e., is either not working in month t+1 or is working but with job tenure of under one month. The second variable is the quit rate, defined as the proportion of people observed on a job in month t who are observed not to be

[7] Starting in 1990, the LFS introduced some modifications to its education questions. The focus changed from measuring years of education to measuring educational attainment. The December 1989 transitions were excluded from our analysis because the numerator and the denominator are not based on the same question. As a robustness check, we repeated our analysis for those with 10 years or less of education, a group for which the effect of the change were minimal (Gower 1993). The results of our regression analysis are essentially the same.

[8] A detailed description of how the data was linked can be found in Appendix A.

[9] We use the period 1 LFS weights in constructing each transition rate.

working in month t+1 and who respond to the question of why they separated from their last job by saying they quit. The third dependent variable is the layoff rate, defined as the proportion of people employed in month t who are not employed in month t+1 and respond that they separated from their previous job due to a layoff. Both the layoff and quit rates miss people who separate from a job and find a new job before the next month. We capture these separations in our fourth variable, job-to-job transitions, which equals the proportion of people working in month t who are working in month t+1 but for a different employer. People on a new job in t+1 are defined as those with job tenure of under one month.[10] The quit, layoff and job-to-job transition rates sum to the separation rate in each period. The fifth dependent variable is the accession rate, which (with some abuse of terminology) we will call the hiring rate, and which we define as the proportion of people who are non-employed in month t that are employed in month t+1. The sixth dependent variable is the proportion of people who say they are not working and not searching for work (not-in-the-labour-force, N, status) in month t who are still not working but are searching for work (unemployed, U, status) in month t+1: the NU flow rate. We also examine the UN flow rate. Finally, we present results where we examine changes in hours of work for individuals that continue with the same employer to see if firms adjusted the work of new hires in this dimension.

Figures 4 and 5 show the Canadian layoff, quit and hiring rate patterns for the 1979-2008 period.[11] As expected, quit and hiring rates are cyclical in nature, while the layoff rate tends to be counter-cyclical. Interestingly, the layoff rate is systematically larger than the quit rates for these workers. Perhaps the most striking feature of these figures is the rapid and substantial increase in the hiring rate in the late 1990s. This coincides with the surge in the employment rate that occurred in Canada over this period.[12]

We present summary statistics for all years combined in Appendix tables. Those statistics match with expectation. In particular, workers with lower levels of initial tenure are more prone to quit or to be laid off, and as such, have a higher separation rate. In addition, younger workers are less likely to continue with the same employer, but they are also more likely to be re-hired.

[10] For the period from 1999 to 2005, the LFS asked job-to-job switchers why they left their first job. Using that data, we find that 68% of workers who transit directly to a new job by the time of the next monthly survey reported that they quit their previous job. Thus, while our job-to-job transition variable largely captures quitters, it clearly also includes a significant number of layoffs.

[11] The monthly rates are averaged over each year.

[12] Campolieti (2011) finds the same patterns when applying Shimer (2007)'s more indirect method of calculating the hiring rate to Canadian data.

3 Empirical Specification and Results

For all of our dependent variables, we use the same estimation specification, as follows:

$$y^g_{p,t} = \alpha^g + \sum_{k=0}^{K} \beta^g_k ln(rmin)_{p,t-12k} + X_{p,t}\gamma^g + \epsilon^g_{p,t} \quad (1)$$

where $y_{p,t}$ is the dependent variable and $ln(rmin)_{p,t}$ is the log of the real minimum wage in province p and period t. The vector of controls, $X_{p,t}$, includes a complete set of provincial dummies, a dummy variable which equals 1 if there was a minimum wage change over the upcoming month, and a full set of time dummies, corresponding to every month of every year. We include the dummy variable for a current minimum wage increase in order to allow us to focus on equilibrium type effects rather than immediate adjustments to a minimum wage change.[13] We present results for the cases where $K = 0$ (i.e. no lags) and $K = 1$ (i.e. using the real minimum wage lagged by one year). We have also estimated specifications with higher lags and report on them where relevant. We use lags specified in years rather than months because we want to examine the possibility of long term adjustments to minimum wage changes. All our estimations are performed using weighted least squares where the weights are the inverse of the number of individuals in the relevant "at risk" group in order to account for the fact that, for example, the number of workers in the province of Prince Edward Island in a month is less than 1.5% percent of the number of workers in Ontario. We employ the g subscript in equation (1) to emphasize that we provide separate estimates for a variety of sub-groups defined by gender, education and age. Once we have obtained results for each of the transition rates, we use them to examine implications for the employment rate in section 3.2.

As is well known, in a case like ours with panel data with few cross-sectional groups, uncorrected OLS standard errors are severely biased downwards (Bertrand et al. (2004)). Hansen (2007) argues that in this situation, a Feasible GLS estimator provides efficient estimates with low size distortion for associated test statistics and is superior to standard clustering approaches, which have poor power properties. All of our reported results are based on a FGLS estimator.[14]

[13] Given that we use relatively high frequency variation, what we capture are likely partial equilibrium effects. The inclusion of lags allows for the possibility that equilibria are approached gradually.

[14] One common concern with FGLS estimators is that the transformations of the variables can lead to very different estimates compared to OLS because different variation is being used. In our case, the FGLS and OLS coefficients are very similar (the latter are available on request). Our estimator is based on an AR3 specification for the error process. We arrived at that process by testing down from higher orders. Given the long length of our time series,

In Table 1, we provide our base results, showing the impact of minimum wages on the separation rate for different initial period tenure levels. For brevity, we do not present the large set of estimated coefficients on the province, time and current period minimum wage change dummy variables. The first panel shows results for males and females combined. The first column presents results not conditioned on initial tenure level and it is apparent that the minimum wage has no impact on separation rates for all workers combined. However, once we examine the impact for workers with less than one year of job tenure, we find a strongly statistically significant coefficient of -.035 on the log real minimum wage variable. This implies that a 10% increase in the real minimum wage leads to a statistically significant 0.0035 decrease in the separation rate for workers with high school or less of education who have been with the same employer less than one year (where a typical one-month separation rate for this group is on the order of 0.07). At first glance, this may appear small but even apparently small changes at the monthly level imply relatively large effects on an annual basis. Thus, the probability a job continues for a year if the monthly separation rate is 0.07 and there is an equal probability of separation in each month is 0.42. If, instead, the monthly separation rate declined to .066 (as the estimate suggests would arise from a 10% increase in the minimum wage), the probability the job continues for a year becomes 0.44 - an approximately 5% increase. What is perhaps even more striking about the estimate, though, is its sign. While basic reasoning from a standard demand and supply model might lead one to expect that separation rates will be higher in high minimum wage jurisdictions, our estimate indicates the opposite effect. As we will see, a negative separation rate effect is consistent with other standard theories of the labour market with some adjustments.

The second set of rows in Table 1 report on a specification including the one year lag of the real minimum wage variable. The coefficient on this lag variable is economically insubstantial and far from statistically significant at any conventional significance level. The implication is that long term adjustments do not reverse the effect of minimum wages in reducing separation rates. In the third and fourth columns in the table we repeat our estimations (with and without the one year lag) for workers whose initial tenure is under six months and workers whose initial tenure is between 6 and 11 months. The results from these exercises reveal that the positive impacts of the

the formula for bias in the AR parameters in a short panel presented in Hansen (2007) implies very little bias in our case (on the order of 0.003 for the first order autocovariance) and so we do not implement corrections for it.

minimum wage for workers of under one year of job tenure displayed in column 2 are entirely due to workers with tenure of under 6 months. The estimated effects are neither economically substantial nor statistically significant for workers with 6 to 11 months of job tenure.

The tenure pattern in Table 1 matches well with regulations on notice requirements for employment termination in Canada. Most workers fall under provincial jurisdiction in terms of labour market regulations. Termination notice varies by length of service in most provinces with many of them having no required notice for short durations. Thus, British Columbia and New Brunswick require no notice for jobs that have so far lasted less than 6 months, and Alberta, PEI, Ontario, Quebec and Saskatchewan require no notice for jobs of less than 3 months duration. For all jurisdictions, the required notification period (when one is required) for jobs of under a year duration are either one or two weeks (Kuhn 1993). Firms are required to pay a lump sum equal to the wages for the notification period if notice is not given. Whether or not these regulations are enough to induce changes in firm layoff behaviour, the law certainly acknowledges what is an effective probationary period during which there is little or no implication from laying off a worker, and the results in Table 1 indicate that minimum wage impacts on terminations occur mainly within this probationary period.

In the second and third panels of Table 1, we repeat all of these estimations separately for men and women. The key result from these estimations is that the effects of minimum wages on separation are nearly identical for men and women. Given this, we will continue with a combined male and female sample for the remainder of the paper.[15]

In Table 2, we present results for three different age groups: ages 15-19, 20-24 and age 25-59.[16] The previous literature has tended to focus on teenagers as a group for whom one expects the minimum wage to be directly relevant. In this data, we observe the strongest impacts for teenagers but only to some extent. We observe significant effects even for the "All Tenure" sample of teenagers

[15]In order to test for gender differences in the minimum wage effect, we re-estimated equation (1) using both male and female separation rates and adding a full set of gender interaction terms, testing for the significance of the interaction effects with F-tests. We repeated the same tests for workers with less than one year of tenure, with 6 to 11 months of tenure and with less than 6 months of tenure. In all cases we could not reject the null hypothesis that the minimum wage effect is the same for men and women (the p-values of all test statistics exceeded 0.7). Gender differences in minimum wage effects are similarly unimportant in the specifications throughout the remainder of the paper with the exception of hiring rates. We show those effects broken down by gender in Table 4.

[16]As one would expect, teenagers tend to have shorter tenured jobs; 43% of teenagers have been with their present employer for less than 6 months. For older adults, longer tenured jobs are the norm; 82% of older workers have job tenure of one year or more, whereas only 10.5% are low tenured (i.e. less than 6 months of job tenure). See Appendix Table A.3 for more detail.

compared to essentially zero effects for "All Tenure" for other ages. Interestingly, though, the effects for workers with under 6 months of tenure for all age groups are substantial and similar in size to that observed for teenagers. This may arise because low educated workers who are starting a new job tend to be low wage workers (and thus workers for whom the minimum wage is relevant) regardless of their age.[17] The more substantial "All Tenure" group effect for teenagers may then just reflect that a much larger proportion of teenagers has under a year of tenure.

Table 3 contains the main results in our paper: the impact of minimum wages separately for quits and layoffs. In the first panel, we present results for quits.[18] For workers, with less than 1 year of job tenure, the specification not including any lags of the minimum wage reveals a negative estimated effect of the minimum wage on quits. This fits with our Table 1 results but the estimated effect is only one-tenth the size of what we observed there and is not statistically significant at any conventional level. Interestingly, when we allow for a lagged effect, the estimated coefficient on the one year lag of the minimum wage is larger and statistically significant (though still less than 1/3 the size of the overall separation effect).[19] Moreover, in contrast to our overall results, the largest and most significant effects are for workers with 6 to 11 months of tenure.

The second panel of Table 3 contains the results using the layoff rate as the dependent variable. When no lags are included in the estimation, the minimum wage coefficient is negative, highly statistically significant and approximately 3/4 of the size of the effect on the overall separation rate. When a single lag is included, the current period effect changes very little and the lagged value coefficients are economically small and not statistically significant. This suggests no diminishing of minimum wage effects on layoff rates over the longer run.

The third panel of Table 3 contains the job-to-job transition rate effects. Recall from footnote 10, that evidence from a subset of our sample years indicates that 68% of job-to-job transitions are quits. The weighted averages of the estimated quit and lay-off coefficients using these relative shares as weights are -.011 for the under 1 year tenure jobs and -.013 for the under 6 month jobs. These are close to the actual estimated coefficients for the job-to-job transitions (-.014 and

[17]The average wage for workers in our 2008 sample that have under 6 months of job tenure was $14.05 while for all other levels of tenure it was $18.62.

[18]Given that minimum wage effects tended to be statistically insignificant and economically insubstantial for job tenure over a year in the previous tables, we mainly focus on results for workers with under one year of job tenure for the remainder of the paper.

[19]When we allow a second year lag in the real minimum wage, the coefficient on this additional lag is not statistically significant and the pattern presented in the table remains.

-.017, respectively), suggesting that job-to-job transitions are not particularly special relative to separations where a job has not been found by the time of the next survey.

We can use these estimated coefficients to assess the relative importance of the different types of separations. In simple estimation with only the minimum wage variable as a covariate, the quit, layoff and job-to-job transition coefficients should add up to the estimated coefficient for total separations. However, because the other included covariates (time and province effects) are not restricted to be the same across the regressions, the adding up is not exact. Using the sum of the estimated quit, layoffs and job-to-job transition coefficients as our base, the effect of minimum wage changes on layoffs in jobs with under 1 year of tenure accounts for 59% of the effect on total separations and the effect in reducing job-to-job transitions accounts for approximately 33%. If we aportion 32% of the job-to-job transition effect to layoffs (because 32% of job-to-job transitions are layoffs) then the effect of minimum wage increases in reducing layoffs accounts for nearly 70% of the overall effect of the increases in reducing separations, with quit effects accounting for the other 30%. For jobs with under 6 months of tenure, layoff effects account for 72% of the total separation effects after dividing up the job-to-job transition effects. We argue, based on these patterns, any attempt to understand the impact of minimum wages on the separation rate should focus on layoffs.

In Table 4, we examine results for other employment related outcomes. In the top panel, we examine the impact of minimum wages on hiring rates for various demographic groups. The both-genders- combined results in the first column indicate that increasing the minimum wage reduces hiring rates by an economically substantial and statistically significant amount. The second set of rows contains results including the lagged minimum wage. While neither minimum wage coefficient is statistically significant, the fact that both are of similar size suggests that full hiring rate effects may phase in gradually. In contrast to the separation rate effects, there is a gender differential in hiring rate effects, with the effects being much smaller for females. The negative hiring rate effect is much larger for teenagers than for workers of other ages. This is in contrast to the relatively small age differentials in separation effects and suggests that minimum wages will have more substantial effects on employment levels for teenagers.

The second panel contains the estimated impact of a change in the real minimum wage on the proportion of not-in-the-labour-force (N) individuals who transit to being unemployed (U), i.e., the NU rate, in the left column and the UN rate in the right column. With or without the lagged minimum wage, we find that increases in the real minimum wage imply statistically significant declines in the NU rate. For UN rates, we find the opposite effect, but only the estimated coefficient on the lag term is statistically significant. These results suggest that those near the margin of entering or leaving the labour force assess the negative effects on hiring from a minimum wage as outweighing any positive effects in terms of higher wages and longer job tenure. This may imply that, at least for workers near this margin, minimum wage increases are welfare decreasing.

In the lower panel of Table 4, we show estimates of impacts on average weekly hours of work for workers with different levels of tenure. One might hypothesize that minimum wage impacts will show up to some extent in reductions in weekly hours rather than in employment changes. In fact, the table indicates that there is no evidence of an impact on hours in higher minimum wage regimes. This is true regardless of the individual's tenure level.

3.1 Robustness

Table 5 contains results from a set of robustness exercises. Given our results so far, we focus on lay-off rates and overall separation rates for jobs with less than 6 months of job tenure for males and females combined, but the outcomes of these exercises are similar for other types of separations and for jobs lasting under 1 year. We present the estimate from our base specification (from Tables 1 and 3) in the first column for comparison.

It is common in the minimum wage literature to capture minimum wage changes using a variable defined as the ratio of the nominal minimum wage to a relevant comparable wage such as the average manufacturing wage. We have, instead, chosen to work with the real minimum wage out of concern about potential endogeneity of the comparison wage in the denominator of a ratio variable. Nonetheless, in the second column we present results using the log of the ratio of the minimum wage (times 40) to the average weekly wage for males with a high school or less education. The estimated minimum wage effects are very similar to those using the real minimum wage, with the effect being slightly smaller for all separations and larger for layoffs.

In the third column, we present a specification designed to check whether our results are being driven by reductions in the real minimum wage due to inflation rather than the actual increases in the nominal minimum wage. If that were true then our results could be sensitive to the specific deflator used. To check this, we instrument for the real minimum wage using the nominal minimum wage. The instrumented coefficients are again negative and slightly larger than the non-instrumented, implying that our results are being driven by the legislated minimum wage changes.

We were also concerned that anticipation of minimum wage changes by firms may imply that our research design is not clean. In our data, we can observe situations where anticipation is clearly possible: minimum wage increases that are announced in advance of actual implementation. To see whether potential anticipation alters our results,in the fourth column of Table 5, we present results based on a sample where we omit all observations between announcement and actual enactment dates. Again, the estimated coefficients are slightly larger than the base case and there is no change in conclusions.

In the fifth column, we present a falsification exercise where we estimate our specification for males, aged 25 to 54 with a BA; a group for whom we would expect to find no minimum wage effects. The results indicate this is indeed the case. For total separations, the estimated coefficient is less than half that for the base case and no where near statistical significance. For layoffs, the estimated coefficient is -0.0001. Thus, at least by this measure, we do not appear to be inadvertently picking up other factors with the minimum wage variable.

Finally, we were concerned that the minimum wage changes may not be exogenous. In particular, one might find a negative correlation of minimum wage changes with layoff rates if governments tend to increase minimum wages during good economic times. We address this concern in two ways. In the sixth column of Table 5, we present results from our standard specification omitting labour market downturn periods (periods with unemployment rates over 8%) in order to eliminate cyclical variation as a source of identification.[20] The results using these periods are very similar to the base case results. In the seventh column, we present results from an IV specification where we instrument for the minimum wage using political variables. In particular, we use a dummy for whether the governing party in the month and province is left wing and a dummy for whether the governing party is right wing. The omitted category corresponds to a centrist party. We also included the average values for these variables across all other provinces in the region, which is a valid instrument within a model in which provincial policy parameters are set partly in relation to the parameter values in other provinces around them (see Green and Harrison (2011)). The IV estimated coefficients are again negative and much larger than those in the base case (implausibly

[20]The included months are: 1979m1 to 1981m12, 1986m9 to 1990m11, and 1997m1 to 2007m12.

so). Thus, there is no evidence that endogeneity stemming either from cyclically sensitive minimum wage changes or other causes is driving our key result.

3.2 Employment Rate Implications

A natural question is the implication for the overall employment rate of the changes in separation and hiring rates induced by a change in the minimum wage. To answer that, we can consider two regimes - one with average hiring, layoff and quit rates for Ontario for the year 2007 and one where we use the estimated effects from Tables 3 and 4 in conjunction with a 10% real increase in the minimum wage to calculate new, counterfactual rates of separation and hiring. We then use the rates from both scenarios to construct the average level of the employment rate in each regime and, from that, the impact of the minimum wage increase on the employment rate.

We calculate the employment rate as the equilibrium rate under the assumption that flows into employment equal flows out. Thus, the equilibrium rate equals hr/(hr + sr), where hr is the hiring rate and sr is the separation rate. We compute the separation rate as the sum of the layoff rate plus the quit rate, noting that direct job-to-job transitions are not relevant for this exercise since we only need flows out of employment. For all ages, the implied employment rate in the base scenario is 69%. This compares favourably with the actual employment rate of 66.6% for high school dropouts and graduates aged 15 to 54 (both sexes) in Ontario in 2007. The calculated minimum wage impact implies that a 10% increase in the real minimum wage generates a 0.76% decline in the employment rate with an associated standard error of 0.50. In comparison, for teenagers, the calculated minimum wage effect is -1.7% with a standard error of 0.9.[21]

To check on the comparison of these calculated minimum wage impacts with those from a more typical approach, we implemented a standard specification using Canadian provincial data for our same sample period (1979 to 2008). Specifically, we regressed age specific employment rates on the real minimum wage, a lag of the minimum wage variable, and a complete set of year and province dummies. For the regressions for teenagers we also included the proportion of the provincial population who were teenagers and the adult male unemployment rate. This specification has been used in numerous previous papers (see Neumark and Wascher (2008) for a recent comprehensive

[21] The implied base equilibrium employment rate for teenagers is 62%, which is higher than the average employment rate of 50.7% for high school grads and dropouts aged 15 to 24 in Ontario in 2007. The fact that we drop full-time students in our sample but they are included in the published statistics could account for this difference.

survey of the minimum wage literature). The results from these estimations implied that a 10% increase in the minimum age would lead to a statistically significant 2.5% decline in the teenage employment rate and a statistically insignificant 0.5% decline for the overall employment rate. These are in substantial agreement with earlier estimates in the literature and fit well with our calculated impacts on the equilibrium rate. Our results indicate that the small net effect on the overall employment rate reflects offsetting negative effects of a minimum wage increase on hiring and layoff rates. For teenagers, the more substantial negative net effect reflects the fact that, for them, the negative impact on hiring substantially outweighs the negative effect on layoffs.

4 Theoretical Implications

To this point, we have found, examining a substantial amount of data on quits, layoffs and hires, that both separation and hiring rates are lower in higher minimum wage regimes. Importantly, the reduction in separations occurs mainly in the first 6 months of a job and mainly through a reduction in layoffs. In this section, we consider the implications of these results for standard labour market models and discuss how they need to be adjusted to fit with the patterns observed here. Based on these insights, we then describe a partial equilibrium model of firm decision making which captures the main features of the data and provides some intuition for them. Our emphasis is on implications for the demand side of the labour market, and so we do not pursue the interesting result that increases in the minimum wage imply declines in NU flows.

4.1 Implications for Standard Models

The most common model used in analyses of minimum wage effects is the static labour demand model with wages set equal to the value of marginal product in the absence of a minimum wage. In such a model, an increase in the minimum wage will result in (weakly) decreasing employment but the model is necessarily silent on how the employment adjustments take place and has nothing to say about hiring and separation rates before and after a minimum wage change. Thus, a static labour demand is not useful for interpreting our results.

Rebitzer and Taylor (1995) examine minimum wage effects in an efficiency wage model. In such a model, an increase in the ratio of the minimum wage to the average wage in the economy could reduce shirking and, therefore, terminations associated with a worker caught shirking. However, they show that the associated reduction in monitoring costs for firms leads to an increase in hiring, which is the opposite of what we find.

Both Portugal and Cardoso (2006) and Dube et al. (2012) invoke Burdett-Mortensen type models to explain decreases in the separation rate when the minimum wage increases. From our perspective, a key feature of these models is the centrality of on-the-job-search and quits. These are at the heart of the models' explanation for why homogeneous firms can co-exist while paying heterogeneous wages. Firms paying low wages realize high per-period profits while the job is filled but their overall profitability is lowered by the fact that their workers are likely to locate a job paying a higher wage and quit, leaving the firm with a costly, unfilled vacancy. High wage firms have lower per period profits from filled jobs but less time with jobs unfilled. In this context, it is job-to-job transition rates or, possibly, quit rates with short intervening unemployment spells that must be the location of any minimum wage effects. In contrast, our results indicate that the main minimum wage effects operate through layoffs, which are not a part of these models. While theoretical distinctions between quits and layoffs are often hard to make in equilibrium models (since separations are mutually agreed upon), it is difficult to believe that, in responding to a survey, workers would label a separation in which they took up a higher wage offer at another firm as a layoff. Thus, we view our results as not fitting with the channels emphasized in these models.

4.1.1 Minimum Wages in a Mortensen-Pissarides Model

A standard Mortensen-Pissarides model, extended to incorporate endogenous separations (Pissarides (2000), chapter 2), is dynamic in nature and includes separations that look plausibly like layoffs. Given the popularity of this class of models, we will focus the rest of our discussion on them.[22] In the Pissarides model, workers and firms meet according to a matching technology after which, in each instant, there is probability with which they draw a new, match-specific productivity value. We will work with a slightly simplified version of this model in which the match-specific pro-

[22] This is not to say that there may not be other models that could rationalize our results. For example, Acemoglu (2001) investigates a search and bargaining model in which firms can create one of two types of vacancies: lower productivity-low wage jobs (bad jobs) or higher productivity-high wage jobs (good jobs). A minimum wage above the bad job sector wage but below the wage in the good sector will reduce relative profits in the bad job sector and cause a shift in composition toward good jobs. Such a model might be extended to include more investment by firms in workers in good jobs and, with it, lower layoffs similar to the mechanism in Acemoglu and Pischke (1999).

ductivity draw occurs only once - at the end of an initial probationary period. This simplification makes the intuition in the model more transparent and fits with the empirical result that minimum wage effects on transitions occur at low job tenures. The key conclusions are unchanged if we use the full Pissarides model.

With this in mind, consider a matching model in which each operating firm employs one worker. Workers and firms meet according to a matching technology but do not know the ultimate productivity of the match when they first meet. The productivity, x, which is a random variable with associated cumulative distribution function F defined over a range $[\underline{x}, \overline{x}]$, is match specific and is not revealed until after the worker has worked for the firm for a brief probationary period. During the probationary period, workers do not produce anything and are paid the minimum wage, m. Pissarides (2000) shows that the problem has a reservation quality such that matches with productivity draws that are less than an endogenously determined value, R, are terminated. If $x > R$, the match continues and the firm and worker bargain a wage to divide the match specific surplus according to a Nash bargaining rule. Both firm profitability and wages, w(x), are increasing in x. Given this, we can define a productivity value, x_m such that $w(x_m) = m$, where m is the minimum wage. For values of x between R and x_m, matches continue but are paid the minimum wage and involve lower profits for firms than non-minimum wage jobs.[23] Firms also pay a flow cost of posting a vacancy, c, and firms and workers face a common discount rate, r.

The key intuition in the model can be obtained by examining the firm Bellman equations, with the equation corresponding to a filled vacancy with a productivity draw, $x > x_m$ being given by,

$$rJ_{nm}(x) = x - w(x) + \lambda(V - J_{nm}(x)) \tag{2}$$

where J_{nm} is the value to the firm of a filled, non-minimum wage paying job, λ is the exogenous probability with which matches end, and V is the value of an unfilled vacancy. The equation for a filled vacancy with $R < x \leq x_m$ is,

$$rJ_m(x) = x - m + \lambda(V - J_m(x)) \tag{3}$$

[23]This follows Flinn (2006), which analyses minimum wage effects in a matching model.

Assuming free entry of firms, V = 0. Given that the value of a filled vacancy is increasing in x, the threshold productivity value at which a firm is just indifferent about continuing with a match will occur on a minimum wage paying job. Thus, R is defined from the condition $J_m(R) = 0$. Using this in equation (3), we get, R=m. Thus, if m increases, the threshold productivity increases and, as a result, layoffs increase. Thus, a canonical search and matching model generates the opposite of our main empirical finding.[24]

One plausible way to extend the standard model so that it can generate decreases in the layoff rate with increases in m is to assume that firms can screen workers before the probationary period with an imperfect measure of productivity. We present an exposition of such a model in Appendix B1. In this model, workers and firms again draw a productivity value, x, when they first meet but they actually observe $\hat{x} = x + \epsilon$, where ϵ is a draw on a random variable corresponding to observational noise. Then, with instantaneous probability, δ, the true productivity value is revealed (ending the probationary period) and firms make an endogenous layoff decision as before. Firms choose between not hiring workers at first contact (with the risk that they will not take on a worker whose true match productivity is high because of imperfect information) or laying them off later when the true productivity is revealed. The cost of the latter strategy is increasing in the minimum wage since workers are paid the minimum wage during the probationary period. Thus, when minimum wages rise, firms screen out more workers at first contact, implying reductions in both the hiring and layoff rates. This fits with our results. However, because layoffs are still ultimately determined by true productivity, the model implies that the employment rate will decline with an increase in the minimum wage, which does not fit with standard empirical findings of no effect on employment rates for older workers.[25]

An increase in m necessarily implies a reduction in the employment rate in these models because the threshold value, R, is increasing in m for all firms. To generate the result that an increase in

[24]The result that $\frac{\partial R}{\partial m} > 0$ continues to hold in the more complicated Pissarides (2000, chapter 2) model with ongoing productivity draws for a match. In that model, the value function for a filled job includes terms corresponding to the option value of continuing with the match (since there is a possibility of drawing a better productivity without having to pay the costs associated with starting a new vacancy). Since one can show that minimum wage matches are less profitable than non-minimum wage matches, that the proportion of jobs that are minimum wage matches is increasing in m, and that the profitability of minimum wage matches is declining in m, the option value of continuing with the match declines with increases in m. This serves to further increase R when m rises.

[25]Indeed, the negative effect of an increase in m on employment will be larger in this model than in the simpler endogenous job destruction model. The increase in R resulting from the rise in m will lead to more layoffs, as in the simpler model, but increased screening at the hiring stage will mean, in addition, that more workers with $x \geq R$ will also be incorrectly terminated at the hiring stage.

m results in no change in the employment rate because reduced hiring and reduced layoffs offset one another, a model needs to allow for $\frac{\partial R}{\partial m} < 0$, at least for some firms. To understand what this implies, note that one can define R as the level of productivity at which the match surplus is zero, i.e., the productivity that just covers the flow values of worker and firm outside options:

$$R = w^* + rV \tag{4}$$

where, w^* is the worker's reservation wage. So far, we have assumed that free entry implies V=0 and that $m > w^*$, so that R = m. For the model to allow $\frac{\partial R}{\partial m} < 0$ requires that $V \neq 0$ and that $\frac{\partial V}{\partial m} < 0$. We present a model with these features in the next section.[26]

4.2 A Model

In this section, we present a Mortensen-Pissarides model with $V \neq 0$ by introducing firm heterogeneity in the cost of opening vacancies, as in Fonseca et al. (2001) (and Beaudry et al. (2011)). In particular, assume that entrepreneurs draw a value for the fixed cost of creating a vacancy from a cumulative distribution function, K. Entrepreneurs with a fixed cost below V will open a vacancy and search for a worker to fill it. This allows for free entry in the sense that entrepreneurs enter until the marginal entrant has zero expected profits, but the value of a vacancy is not driven to zero by competition. The question is then whether an increase in m could lead to a reduction in V and, hence, a potential negative effect on R.

The matching function determining firm and worker meetings is a constant returns to scale function of labour market tightness, $\theta = \frac{V}{U}$ (V = the number of vacancies and U = the number of unemployed workers). The matching probability implies a probability of a firm's vacancy meeting a worker, $q(\theta)$, and a probability of an unemployed worker meeting a vacancy, $\theta q(\theta)$. There is no on-the-job search. Matches end according to an exogenous probability, λ, but will also be terminated

[26] Inspection of (4) may suggest that we could also get $\frac{\partial R}{\partial m} < 0$ if there were heterogeneity in w^* such that some workers have $w^* > m$ and $\frac{\partial w^*}{\partial m} < 0$. For matches with workers with $w^* \leq m$, R will always equal m, as we have seen, implying that movements in w^* caused by changes in m are irrelevant for determining layoffs. For workers with $w^* > m$, the worker's reservation wage will help determine the relevant R. If firms reduce hiring as m increases, a reduction in w^* could result. Indeed, our NU transition results indicate that the value of search is lowered for workers on the margin of participation. However, if this were true for all workers then it would imply reductions in bargaining power and, hence, wages for above minimum wage workers. This is the opposite of what has been found in papers examining wage spill-over effects (Card and Krueger (1995), Neumark and Wascher (2007)). For this reason, we do not pursue this avenue.

endogenously when x is revealed in some cases. We will present results in a partial equilibrium setting in which firms take θ and worker reservation wages as fixed. In this setting, we will discuss the separation decision as if it is made unilaterally by the firm and call the separations layoffs. In general equilibrium, the decision to dissolve a match is, in fact, jointly agreed upon in most, but not all, situations. Since the main intuition is easier to see in the partial equilibrium setting, we present the more restrictive model here and leave an exposition of general equilibrium results for future work.

The value function corresponding to a vacancy is given by,

$$rV = -c + q_F(\theta)(J_F^e - V) - q(\theta)m \tag{5}$$

where, J_F^e is the expected value of a filled vacancy and we have assumed that a firm has to pay the probationary wage (equal to m) when it meets a worker. It then observes the true value, x, and determines whether to lay-off the worker. Thus, $q_F(\theta)$ is the probability a match is made and continues. That is, $q_F(\theta) = q(\theta) \times Probability$ (the match specific productivity is high enough to warrant continuing).

For reasons we will discuss below, workers are assumed to be heterogeneous with respect to their reservation wages, w_j^*, where j indexes the worker and he or she prefers the unemployed state to being employed with a wage below w_j^*. More specifically, we will assume that workers draw individual values of the flow value of non-employment, b_j, from a distribution with CDF, H, and those then imply heterogeneous reservation wages. Because of this, the bargaining solution, and with it, profits, will vary with the specific worker involved in the match. As before, bargained wages will be increasing in x and there will be set of minimum wage paying jobs with value function given in (3). For matches with higher x draws the value function will be given as in (2) but the wage now needs to be indexed by j since the bargained wage will depend on the specific worker's outside option. The reservation productivity will also vary by worker if $w_j^* > m$:

$$R_j = rV + w_j^* \tag{6}$$

Workers with $w_j^* \leq m$, will face a common reservation productivity value given by,

$$R_m = rV + m \tag{7}$$

In Appendix B, we show that one can write,

$$rV = \frac{r+\lambda}{r+\lambda+q_F(\theta)}(-c - q(\theta)m) + \tag{8}$$

$$\frac{q_F(\theta)}{r+\lambda+q_F(\theta)}\{\int_{\underline{b}}^{b_m} E(x-m \mid R_m < x < x_m(s)) \times \frac{F(x_m(s)) - F(R_m)}{1-F(R_m)}$$

$$+E(x - w(x,s) \mid x > x_m(s))\frac{1-F(x_m(s))}{1-F(R_m)}h(s)ds\} +$$

$$\frac{q_F(\theta)}{r+\lambda+q_F(\theta)}\int_{b_m}^{\bar{b}} E(x - w(x,s) \mid x > x_m(s))h(s)ds$$

where $w(x,b)$ is the wage that would be bargained in a match with productivity, x, between a firm and a worker with flow value of unemployment, b. The first term on the right hand side of (8) corresponds to the discounted cost of filling the vacancy and the following three lines correspond to expected flow profits from a match.

To understand the expected flow profits, note that it is relevant to divide workers into two types: those with a reservation wage below the minimum wage and those with a reservation wage above the minimum wage. More specifically, we can define b_m as the value of the flow value of unemployment for a worker such that their reservation wage, $w_j^* = m$ (given current market conditions). In addition, we now define $x_m(b)$ as the productivity level such that a firm and a worker with flow value of unemployment, b, would just bargain a wage equal to m. Thus, for $R_m \leq x < x_m(b)$ the match will continue with the worker receiving the minimum wage. From this one can see that the second and third lines of (8) is the expected profits from low reservation wage workers, with the first part corresponding to profits from matches paid the minimum wage and the second corresponding to matches where x is high enough that a wage above m is paid. The fourth line of (8) corresponds to expected profits from workers with reservation wages above m.

Inspection of 8) reveals that, given that the value of m is assumed not to affect productivity draws, $\frac{\partial V}{\partial m}$ will depend crucially on the derivative of bargained wages ($w(x, b)$) with respect to m. In the appendix, we show that we can write:

$$w(x, b_j) = w_j^* + \beta(x - w_j^* - rV) \quad (9)$$

where β is the bargaining parameter determining the share of the match specific surplus going to workers rather than firms.

Recalling that firms take worker reservation wages as fixed, we then get:

$$\frac{\partial w_j}{\partial m} = -r\beta \frac{\partial V}{\partial m} |_{R_m} \quad (10)$$

where we now write w_j instead of $w(x, b_j)$ since this derivative is not a function of x.[27] Using (10), we get:

$$\frac{\partial rV}{\partial m} |_{R_m} = -\frac{q(\theta)(r + \lambda)}{r + \lambda + (1 - \beta P_{Nm})q_F(\theta)} - \frac{q_F(\theta)}{r + \lambda + (1 - \beta P_{Nm})q_F(\theta)} P_m \quad (11)$$

where, P_m is the probability a new match will end up being a minimum wage paying match and P_{Nm} is the probability a new match will end up with a wage above the minimum wage.

The derivative in (11) is negative. In particular, the first term on the right hand side corresponds to the increased hiring costs associated with paying workers the minimum wage during the probationary period when m rises. The second term corresponds to the fact that, with probability $q_F(\theta)P_m$, a vacancy will ultimately be filled with a minimum wage match and the profits of such matches are directly declining in the minimum wage.[28]

Finally, returning to (6), the derivative $\frac{\partial R_j}{\partial m} = \frac{\partial rV}{\partial m} |_{R_m}$ for workers with a reservation wage above m. Thus, for these workers, an increase in m leads to a reduction in layoffs because the firm's outside option has declined. Essentially, having already matched with a worker and paid the probationary period wages, the higher is the minimum wage, the more likely is the firm to want to maintain the current match than go back, re-pay the probationary period costs and potentially

[27] Note that terms involving $\frac{\partial R_m}{\partial m}$ equal zero by the envelope theorem.

[28] Here, the firm entry specification plays a role. With the specification set out here, the elasticity of supply of entrepreneurs is determined by the shape of the K distribution. If the supply is less than perfectly elastic then V is non-zero and can be affected by m. Beaudry et al. (2011) present evidence that the supply of entrepreneurs, and with it the job creation curve in a standard search model, is relatively inelastic with respect to changes in wage costs.

end up in a lower profit, minimum wage match. On the other hand, for workers with a reservation wage below m,

$$\frac{\partial R_m}{\partial m} = \frac{\partial rV}{\partial m} |_{R_m} +1 \quad (12)$$

Thus, for these workers, the effect of the reduced outside option is offset by the direct increase in cost from paying m in the marginal matches. We would expect that this direct effect would be larger than the indirect effect in (11) and thus that $\frac{\partial R_m}{\partial m} > 0$. Whether the ultimate impact on layoffs is negative or positive then depends on the relative numbers of matches of each type and is an empirical matter. On the other hand, the decline in the value of vacancies has the straightforward implication that firms open fewer vacancies, resulting in a lower hiring rate. The net implication for the employment rate is, again, an empirical matter.[29]

4.3 Empirical Implications

Having arrived at a model that can rationalize our main empirical results, we now turn to examining further empirical implications of the model. The mapping from the model to empirical implications is based on the fact that in the model, the probability of a layoff is the probability the match specific probability, x, is less than the relevant reservation value (either R_j or R_m). The empirical specification in (1) can then be seen as a linearization of these probabilities using the expressions (6) and (7) for the reservation values. From these expressions, we would want to control for factors that move either V or w_j^* and are correlated with m. Our empirical specification controls for such factors that are province-specific and time-invariant (such as provinces which perennially pursue left wing policies of generous social assistance benefits and high minimum wages) and that take the form of common time effects. The primary empirical implication from the model is then that, if workers with reservation wages above m are more numerous than those with reservation wages below m then layoffs should decline with increases in m and that this layoff effect should be concentrated in the first months of a job when match productivity is revealed.

Of course, the model was selected in order to match these patterns so this does not constitute a

[29]It is at this point that the need to assume heterogeneity in reservation wages across workers becomes apparent. If all workers had a reservation wage above m then there would be no minimum wage effects. Alternatively, if all workers had a minimum wage below m then the direct effect of the minimum wage change would dominate and we would return to the problem with the standard model. Note that we need both worker and firm heterogeneity since without firm heterogeneity, free entry would push V to zero, eliminating any indirect effects of m on R_j.

test of the theory. However, the model has other, testable implications. A first implication comes from (11), which shows the derivative of the firm's outside option with respect to the minimum wage. In situations where a new vacancy is more likely to meet with a minimum wage match, a higher minimum wage will have a larger negative effect on π_v, i.e., higher minimum wages should lead to a larger decline in layoffs. We examine this using an assumption that high school drop-outs are more likely to be minimum wage workers than high school graduates.[30] In particular, we create a variable equalling the proportion of individuals aged 15 to 34 with a high school or less education in a given province and month who are drop-outs. We estimate a specification with this variable entering as a covariate on its own and interacted with the minimum wage variable. In the separation equation for jobs with under 6 months tenure, the coefficient on the log of the real minimum wage in this specification is -.029 (with a standard error of .016) and that coefficient on the interaction term is -.077 (standard error of .079). While the latter effect is not well defined, it does go in the predicted direction.[31]

The intuition behind the main result of the model suggests a second potential empirical implication. In particular, minimum wage increases can reduce layoffs in the model because of firm expectations about the value of future vacancies. It seems reasonable to predict that this effect would be lessened when there is inflation and nominal minimum wages are not adjusted to offset it. In that case, predicted future real minimum wage effects would be lessened. To determine whether this logic might be true, we need to extend the base model to include inflation. In particular - in order to highlight the role of minimum wages - we construct an extension in which firms and workers are assumed to flexibly adjust their bargained wage to keep it constant in real terms. However, nominal minimum wages are not adjusted and, so, are eroded by inflation. Specifically, we consider m in the model to be generated as, $m_t = m_0 e^{-\rho t}$, where m_t is the real minimum wage in period t, m_0 is the initial nominal minimum wage and ρ is the inflation rate.

With a changing minimum wage, the values for posted and filled vacancies will now include

[30] This does not fit strictly within the model as presented since it does not include observable skill characteristics. But an extension to include such characteristics is straightforward and shows that workers with low observed skills are more likely to be minimum wage workers.

[31] At first glance, it may appear that we can identify low reservation wage workers as individuals being paid the minimum wage. However, we cannot use observed wages in the data to separate workers of different types because, within the model, worker type revelation and the layoff decision happen at the same time. When we observe a worker's wage, we already know he or she is not being laid off. In addition, some minimum wage earners will be workers in the probationary period.

terms capturing expected changes in those values (or "capital gains" from those assets). That is,(30) and (2) become,

$$rV = -c + q_F(\theta)(J_F^e - V) - q(\theta)m + \dot{V} \tag{13}$$

and

$$rJ_{nm}(x, b_j) = x - w(x, b_j) + \lambda(V - J_{nm}(x, b_j)) + J_{nm}(\dot{x}, b_j) \tag{14}$$

where, $J_{nm}(x, b_j$ is the value of a filled, non-minimum wage job with productivity draw, x, and filled by worker j. The last terms in (13) and (14) are the "capital gains" terms. In the third section of Appendix B, we show that inflation does have the effect of making $\frac{\partial V}{\partial m}$ and, hence, $\frac{\partial R_j}{\partial m}$ less negative. Thus, this extension serves to emphasize the crucial role played by the impact of the minimum wage on the expected future value of vacancies in the model and to provide a means to test that role. To the extent the data shows that the negative impact of the minimum wage on layoffs is lessened with inflation, this fits with the expectational channel emphasized in the model. This is particularly the case since we use the real minimum wage in our estimates, implying that the direct role of inflation in changing the value of the minimum wage in the current period is already accounted for.

We test this implication by interacting provincial level inflation with the real minimum wage variable. According to our theory, the coefficient on that interaction should be positive: in high inflation periods the layoff reducing effect of minimum wage increases should be lessened. This implication stems from a combination of match productivity not being immediately revealed (so there are layoffs) and free entry not driving expected profits from a vacancy to zero (so that those profits can vary with inflation). Thus, this implication could arise in a variant of a dynamic labour demand model that incorporates these features but not in a standard Burdett-Mortensen type model. In the latter model, separations arise from workers quitting to take a higher wage offer. Even in variants where the current firm can make counter-offers, this will depend on the productivities of the current and offering firms, not on expectations about future values of the real minimum wage.

Table 6 contains the results from specifications involving the inflation interaction. The inflation rates used in the estimation are province specific. Given the inclusion of a full set of time dummies, the effect of this interaction is identified from relative differences in inflation within provinces over time. We include the inflation rate both directly and in interaction with the real minimum wage variable. The inflation rate on its own results in statistically significant declines in the separation and layoff rates. Its impact on the quit rate is smaller than for the layoff rate and generally not statistically significantly different from zero at conventional significance levels. These results fit broadly with a situation where (in contrast to our model) firms and workers bargain nominal rather than real wages and higher inflation allows for declines in real wages that can result in firms not laying off workers. The interaction term implies a strongly statistically significant effect of inflation in the direction of mitigating the negative effect of the real minimum wages on the separation rate. Again, this effect arises mainly through layoffs, with higher inflation periods being associated with a less negative impact of the real minimum wage on the layoff rate. The fact that the inflation impact occurs through layoffs fits with the model. As discussed earlier, there is no apparent reason why quits should respond to the inflation regime in this way in models such as Burdett-Mortensen models which emphasize quits. In that sense, these results lend more support to models that emphasize layoffs rather than quits as the channel through which minimum wages reduce the separation rate.

The results in Table 6 imply very substantial impacts of the real minimum wage on layoff rates in low inflation regimes. Thus, for example, with an inflation rate of 2% (close to Canada's average over the last decade), a 10% increase in the real minimum wage implies a 4.3% decline in the separation rate due to layoffs for jobs with tenure under 1 year. This compares to the overall average of a 2.5% decline shown in Table 1 and to a zero effect if the inflation rate reaches approximately 8%. Thus, the results in this table both provide support for the model and imply that the impact of minimum wages on layoffs are even more substantial in recent times than what is shown in the earlier tables.

5 Conclusion

In this paper, we investigate whether and how employment transitions differ in high versus low minimum wage regimes. We do this using data from the Canadian Labour Force Survey which has a consistent question on job tenure throughout our sample period (1979-2008). This allows us to take advantage of the fact that the minimum wage is set at the provincial level in Canada, resulting in over 140 minimum wage changes in our period. We focus on low educated workers throughout. Working with this data, we find that higher real minimum wage regimes are associated with lower job separation rates and lower hiring rates, with both effects being economically substantial and statistically significant. Our most important result is that the reduction in separation rates is driven mainly by a reduction in layoffs rather than quits. We also find that this reduction occurs mainly in the first six months of a job and that the size of the effect at the outset of a job is similar across age groups (including teenagers) and between genders. Finally, we find that flows from out of the labour force into unemployment are reduced when the minimum wage rises.

In the fourth section of the paper, we discuss standard labour market models in light of these results, arguing that Burdett-Mortensen type search models, efficiency wage models, and the canonical form of the Mortensen-Pissarides model do not imply these patterns. We then present a modified form of a Mortensen-Pissarides model with endogenous job destruction that can explain the patterns. That model implies that firms operating in the low skilled labour market reduce layoffs because their expected profits from terminating the current match and starting a new one are lower when the minimum wage is higher. This model has an implication for differences in the minimum wage effect with the inflation rate which we find is supported in our data. Once those inflation effects are taken into account, we find that a 10% increase in the real minimum wage when the inflation rate is 2% (which is near the rate for Canada for the last decade) implies a decline in separations occurring through layoffs of 4.3%.

Taken as a whole, these results imply that a higher minimum wage regime is associated with significantly lower hiring rates and lower layoff rates, particularly in the first six months of a job. For the workforce as a whole, these effects almost exactly offset one another, resulting in no net impact on the employment rate. This fits with standard estimations of the impact of minimum wage changes on the overall employment rate for all age groups. Policy makers then face a choice between a high minimum wage regime where workers take longer to find a job but have greater job stability once they match with a firm versus a low minimum wage regime where workers move more quickly through both unemployment and employment spells. Based on this, the key question becomes which regime is associated with higher welfare. The answer to that will depend in part on worker preferences about job stability versus being unemployed. It will also depend on whether greater job stability is associated with greater investment in firm specific human capital. The ultimate welfare impact is beyond the scope of this paper, but the similarity of the estimated impacts across age and gender groups imply that these welfare implications are important for the entire spectrum of low skilled workers. In contrast, estimations focusing just on the net employment rate impact (and ignoring impacts on the underlying gross transition rates) would lead one to conclude that minimum wages have little impact on most workers older than teenagers.

Appendix A

In this appendix we provide more details information on the LFS, and in particular, the construction of our mini panels.

The LFS has a rotating panel design where households remain in the sample for six consecutive months. Every month 1/6 of the sample is replaced by households in a similar area. Although it has panel features, it is not a panel data set per se; the LFS is officially designed to produce cross sectional samples. As such, it follows dwellings, and not individuals. If an individual changes dwelling, he is out of the reach of the survey.

The LFS also does not have a single person identifier variable. Fortunately, we can uniquely identify individual across monthly files using a combination of variables—all of which are provided by the LFS. Changes over time in geographical identifiers (e.g. EI regions) have meant that different identifying variables must be used for different periods. We provide both a short description of the variables and also its name (in capital letters) as identified in the LFS codebook. For the 1976 to 1983 period, one must use the month, the regional office (REGOFF), the unique household identifier within a regional office (DOCKET), and the unique person identifier within a household (LINE) variables. For 1984 to 1986, one must rely on the month, the economic regions (ERTAB), the census metropolitan areas and urban centres (CMATAB), the REGOFF, the DOCKET, and the LINE variables. For 1987 to 1995, it is the month, the ERTAB, the CMATAB, the unemployment insurance region (UIRTAB), the REGOFF, the DOCKET, and the LINE variables. Finally, for 1996 onwards one must use the month, the one-digit province code (PROV1), the pseudo UIC regions (PSEUDOUI), the regional strata (FRAME), the super-stratum (STRAFRAM), the sample design type (TYPE), the first-stage sampling unit (CLUST), the rotation number (ROTATION), the number assigned to dwellings within a cluster (LISTLINE), the multiple dwelling code for structures that have more than one dwelling (MULT), and the LINE variables.

We dropped individuals that had incompatible tenure spells across the two periods of the panel. For an individual that worked in period 1, she must have one more month of tenure in period 2 (i.e. continued with the same employer), one month of tenure in period 2 (i.e. started a new job), or no job tenure in period 2 (i.e. is out of work). Finally we also dropped that transitioned to self-employment.

Appendix B: Theoretical Models (For Online Publication Only)

B1 Model with Two Stage Screening

In this subsection, we derive the implications for transition rates in a model in which firms can screen workers at two different points. As in the text, firms and workers meet according to a matching technology. Once they meet, a match specific productivity, x, is drawn from a distribution with cumulative distribution function, F(x) defined over the range $[\underline{x}, \overline{x}]$. However, neither the worker nor the firm see the actual value of x until after a probationary period. At the outset, they instead observe the productivity value with error. That is, they observe

$$\hat{x} = x + \epsilon \tag{15}$$

where ϵ is a mean zero error, independent of x, with a known distribution, G, defined over a range $[\underline{\epsilon}, \overline{\epsilon}]$.

Given that the worker does not produce anything but is paid the minimum wage, m, during the probationary period, we can define the Bellman equation corresponding to a new hire as,

$$rJ_I(\hat{x}) = -m + \delta(E(J_F|\hat{x}) - J_I(\hat{x})) \tag{16}$$

where, δ is the instantaneous probability that the true productivity is revealed (thus ending the probationary period) and $E(J_F|\hat{x})$ is the expected value of the job conditional on the initial observed value, $\hat{x}$. Note that J_F will equal either $J_m(x)$ or $J_{nm}(x)$ depending on the x value. As before, we assume that free entry implies that the value of an unfilled vacancy, V, equals 0.

Since higher values of $\hat{x}$ imply higher expected values of x, J_I is increasing in $\hat{x}$ and we can define a threshold value, $\hat{R}$, such that firms do not hire workers they meet when $\hat{x} < \hat{R}$. Thus, firms can do an initial screen of workers they meet based on an imperfect measure of productivity. For sufficiently low values of that measure, they decide it is not worth incurring the cost of hiring them for the probationary period. At the end of the probationary period, the true value of productivity is revealed and, as before, workers are laid off if x<R. Since the second part of the problem (after x is revealed) is unchanged, it is again the case that R = m. However, in this model, it is possible that an increase in m will lead to a decrease in the layoff rate to the extent that more workers were screened out at the initial meeting point.

Given free entry and (16), $\hat{R}$ is implicitly defined by

$$\delta E(J_F|\hat{x} = \hat{R}) = m \tag{17}$$

We are interested in establishing the sign of $\frac{\partial \hat{R}}{\partial m}$. We can do this by differentiating (17) with respect to m:

$$\delta \frac{\partial E(J_F|\hat{x} = \hat{R})}{\partial m} - 1 = 0 \tag{18}$$

We can write,

$$E(J_F(\hat{R}) = \frac{1}{r+\lambda} \int_{\hat{R}-x_m}^{\hat{R}-R} (\hat{R} - s - m) g(s) ds + \frac{1}{r+\lambda} \int_{\hat{R}-\overline{x}}^{\hat{R}-x_m} (\hat{R} - s - w(\hat{R} - s) g(s) ds \tag{19}$$

To work with this further, we will need an expression for the wage that would be bargained on a job with productivity, x. We get this from the Nash bargaining equation,

$$W_{nm}(x) - U = \beta(J_{nm}(x) + W_{nm}(x) - U) \tag{20}$$

where, W_{nm} is the value of a non-minimum wage paid job to a worker, U is the value of unemployment, β is the parameter determining the division of the surplus, and we have used V=0.

We can write,

$$W_{nm}(x) = \frac{1}{r+\lambda}(w(x) + \lambda U) \tag{21}$$

and,

$$J_{nm}(x) = \frac{1}{r+\lambda}(x - w(x)) \tag{22}$$

Using (21) and (22) in (20) and assuming that firms take the value of worker outside options (rU) as given, the wage is,

$$w(x) = \beta x + (1-\beta) rU \tag{23}$$

Given this, x_m, the value of x such that the bargained wage just equals m, is given by,

$$x_m = \frac{1}{\beta} m - \frac{1-\beta}{\beta} rU \tag{24}$$

Using these expressions, the result that R=m, and assuming that $(\hat{R} - \overline{x}) > \underline{\epsilon}$, we can obtain,

$$\frac{\partial E(J_F|\hat{x} = \hat{R})}{\partial m} = \frac{1}{r+\lambda} \{ [\int_{\hat{R}-x_m}^{\hat{R}-R} g(s) ds + (1-\beta) \int_{\hat{R}-\overline{x}}^{\hat{R}-x_m} g(s) ds - (\overline{x} - w(\overline{x})) g(\hat{R} - \overline{x})] \frac{\partial \hat{R}}{\partial m} - \int_{\hat{R}-x_m}^{\hat{R}-R} g(s) ds \} \tag{25}$$

Using this in (18), we get

$$\frac{\delta}{r+\lambda}[\int_{\hat{R}-x_m}^{\hat{R}-R} g(s)ds+(1-\beta)\int_{\hat{R}-\overline{x}}^{\hat{R}-x_m} g(s)ds-(\overline{x}-w(\overline{x}))g(\hat{R}-\overline{x})]\frac{\partial\hat{R}}{\partial m}-[1+\frac{\delta}{r+\lambda}\int_{\hat{R}-x_m}^{\hat{R}-R} g(s)ds]=0 \tag{26}$$

Assuming that $g(\hat{R}-\overline{x})$ is (reasonably) small, the term in [ ] multiplying $\frac{\partial\hat{R}}{\partial m}$ is positive. Then, since the term after the negative sign is also positive, $\frac{\partial\hat{R}}{\partial m}$ must be positive. An increase in m leads to a higher cut-off threshold for $\hat{x}$.

The rate at which unemployed workers move into jobs (what we call the hiring rate) in this model is given by

$$h=\theta q(\theta)\int_{\underline{x}}^{\overline{x}}\int_{\hat{R}-s}^{\overline{\epsilon}} g(y)f(s)dyds \tag{27}$$

Thus, working in partial equilibrium, with θ taken as fixed, $\frac{\partial h}{\partial m}<0$ since $\frac{\partial\hat{R}}{\partial m}>0$: an increase in the minimum wage lowers the hiring rate.

The impact of an increase in m on the layoff rate is more complicated. The layoff rate is given by,

$$l=\frac{\int_{\underline{x}}^{R}\int_{\hat{R}-s}^{\overline{\epsilon}} g(y)f(s)dyds}{\int_{\underline{x}}^{\overline{x}}\int_{\hat{R}-s}^{\overline{\epsilon}} g(y)f(s)dyds} \tag{28}$$

That is, the probability of a layoff is the conditional probability that x<R given that $\hat{x}\geq\hat{R}$.

If we call the numerator in (28), N, and the denominator, D, then we can write:

$$\frac{\partial l}{\partial m}=[D*f(R)\int_{\hat{R}-R}^{\overline{\epsilon}} g(s)ds]\frac{\partial R}{\partial m}+[N*\int_{\underline{x}}^{\overline{x}} g(\hat{R}-y)f(y)dy-D*\int_{\underline{x}}^{R} g(\hat{R}-y)f(y)dy]\frac{\partial\hat{R}}{\partial m} \tag{29}$$

The coefficient multiplying $\frac{\partial R}{\partial m}$ is positive since an increase in the threshold level, R, unambiguously increases layoffs. The coefficient on $\frac{\partial\hat{R}}{\partial m}$, however, is ambiguous. Recall that the layoff rate is defined as the number of workers laid off as a proportion of the number initially hired. An increase in $\hat{R}$ reduces the number of initially hired workers, implying a mechanical increase in the layoff rate. However, it also implies greater selection of workers at the first stage which in turn means fewer layoffs are needed when the true productivity is revealed, lowering the layoff rate. Whether this latter effect is large enough to make $\frac{\partial l}{\partial m}<0$ is an empirical issue. Working with a simple example in which both the true productivity and the initial observational error are distributed as uniforms, $\frac{\partial l}{\partial m}$ is more likely to be negative, the smaller is the spread of the error distribution. Moreover, $\frac{\partial\hat{R}}{\partial m}>\frac{\partial R}{\partial m}$ is a necessary condition for $\frac{\partial l}{\partial m}<0$. Since $\frac{\partial R}{\partial m}=1$ this means we need $\frac{\partial\hat{R}}{\partial m}>1$. An examination of (26) reveals that this is possible as long as δ (the instantaneous probability that the true productivity is revealed) is not substantially larger than $r+\lambda$ (the rate at which the flow from a filled job is discounted). Since it is not clear whether the rate of job destruction is larger or smaller than the rate of revelation of true productivity, this again implies that the outcome is an empirical matter.

The model has one clear prediction: an increase in m implies a reduction in the employment rate. As in the simpler model, when m increases, R increases and more workers are laid off. While an increase in $\hat{R}$ induced by an increase in m can reduce the proportion of employees who are laid off, this is just because of a reallocation of when the firm severs ties with the worker. An increase in m will actually generate a larger drop in employment in the two stage model than in the simple model since anyone with $x<R$ will still, ultimately, be laid off but there will also be matches with $x\geq R$ that will be incorrectly terminated at the initial stage.

B2 Main Model

B2.1 The Environment

Consider an environment in which there is a fixed set of firms, with each active firm hiring one worker. When a firm matches with a worker, there is a match-specific productivity draw, x, of a mean zero random variable with associated CDF, F. Importantly, the value of x is not revealed until after the firm pays a fixed training cost. We will think of that cost as corresponding to a period in which workers add nothing to output but must be paid the minimum wage, m, though we model it simply as a fixed cost, m. Firms also pay a flow cost of posting a vacancy, c, and firms and workers face a common discount rate, r.

Firms and workers meet according to a matching function which is a function of labour market tightness, $\theta = \frac{V}{U}$ (V = the number of vacancies and U = the number of unemployed workers). The matching probability implies a probability of a firm meeting a worker, $q(\theta)$, and a probability of an unemployed worker meeting a vacancy, $\theta q(\theta)$. There is no on-the-job search. Matches end according to an exogenous probability, λ, but will also be terminated endogenously when x is revealed in some cases.

Workers are heterogenous with respect to their flow value of utility while unemployed, b, where $b \epsilon \left[\underline{b}, \overline{b}\right]$ and has an associated CDF, H.

The Bellman equation corresponding to a vacancy for a firm is

$$rV = -c + q_F(\theta)(J_F^e - V) - q(\theta)m \tag{30}$$

where, J^e is the expected value of a filled vacancy and we have assumed that a firm has to pay the probationary wage (equal to m) when it meets a worker. It then observes the true value, x, and determines whether to lay-off the worker. Thus, $q_F(\theta)$ is the probability a match is made and continues. That is, $q_F(\theta) = q(\theta) \times Probability$ (the match specific productivity is high enough to warrant continuing), where the latter equals $q_F(\theta) = 1(\theta) \int_{\underline{b}}^{\bar{b}} \int_{x^*}^{\infty} dF dH$, and x_j^* is the reservation value of x such that matches with $x < x^*$ are terminated. We will discuss the notion that this problem is characterized by having a reservation property in the next section.

The value of a vacancy filled by worker j and with a realized draw, x, is determined by

$$rJ_j(x) = x - w(x, b_j) + \lambda(V - J_j(x)) \tag{31}$$

where $w(x, b_j)$ is the wage bargained when a firm meets worker j with flow value of nonemployment, b_j. We assume the wage is determined by a Nash bargaining solution to the problem of dividing the surplus from a match. Note that the firm's profits are indexed by j because different workers have different outside options, implying different surpluses from the match.

For workers, the Bellman equation corresponding to unemployment is given by,

$$rU_{uj} = b_j + \theta q(\theta)_F \left[U_{ej}^e - U_{uj}\right] \tag{32}$$

where $U_{uj}$is the value of being unemployed, U_{ej}^e is the expected value from employment, ψ is the probability the worker meets a vacancy, and $\theta q(\theta)_F$ is the probability a worker meets a match that is ultimately completed.

The Bellman equation related to employment is

$$rE_{ej}(x) = w(x, b_j) + \lambda \left[U_{uj} - U_{ej}\right] \tag{33}$$

The total surplus to the match is,

$$S_j(x) = J_j(x) + U_{ej}(x) - V - U_{uj} \tag{34}$$

Since the worker's and firm's outside options are independent of x and the benefits to the match are increasing in x, there exists a value R at which the match surplus is zero. Matches with $x < R$ are terminated.

Using (31) and noting that rU_{uj} defines a reservation wage, w_j^*, for worker j, we can write,

$$rV = R_j - w_j^* \tag{35}$$

That is, if a firm just pays the worker his or her reservation wage, R_j^* is the productivity of the match where flow profits just equals the flow value of the firm terminating the match and opening a new vacancy. Note that the left hand side of (35) does not have a j subscript so any increase in w_j^* (i.e., rU_{uj}) is exactly offset by an increase in R_j

Within this model, the interesting termination activity relates to the endogenous decision not to continue with matches where x is below R_j. Our interest, in particular, is in the impact of the minimum wage, m, on terminations, which reduces to understanding its impact on R_j. Apart from its role in determining the costs of training, the minimum wage is only relevant in bargaining if it is higher than rU_{uj} for at least one worker, and we assume this is the case.

B2.2 Partial Equilibrium Impacts of the Minimum Wage on Terminations

B2.2a No Inflation

We begin by considering effects in partial equilibrium from the firm's point of view. In particular, we assume that market tightness (θ) is constant and that the set of reservation wages for all workers is taken as fixed. We will also assume there is no inflation. In the next subsection we will allow for inflation.

We proceed by substituting an expression for rV in terms of basic parameters into (35). Returning to (23), this involves getting an expression for J^e. To do this in the presence of a minimum wage, we need to consider two types of workers: those whose reservation wage is above and those whose reservation wage is below m. Thus, define $b_m$as the value of b_j such that $rU_{uj} = m$, (given current market conditions). Then, we can write,

$$rJ^e = \int_{\underline{b}}^{b_m} \left[\lambda V + E(x \mid x > R_m) - m\frac{F(x_{ml}) - F(R_m)}{1 - F(R_m)} - E(w \mid x > x_{ml})\frac{1 - F(x_{ml})}{1 - F(R_m)} \right] dH \tag{36}$$

$$+ \int_{b_m}^{\bar{b}} [\lambda V + E(x \mid x > x_{ml}) - E(w \mid x > x_{ml})]\, dH - \lambda J^e$$

where, R_m is the reservation productivity draw that allows a firm to just cover the minimum wage plus rV. Notice that this is the same for all pairs where $rU_{uj} < m$ and implies more layoffs than would arise without a minimum wage. x_{ml} is the value of x such that a firm and worker l would just bargain a wage equal to m. Thus, for $R_m \le x \le x_{ml}$ the match will continue with the worker receiving the minimum wage. x_{ml} is indexed by l since it depends on the worker's value of unemployment.

For workers with $b > b_m$, the minimum wage does not have a direct effect on decisions because no such worker would ever be paid that minimum wage (since it is below his or her reservation wage). R_j in this case is defined in equation (6) in the text. Note that m will still have an indirect effect for matches involving these workers since it will affect V_{uj} and V.

Substituting (36) into (23) and rearranging, we get:

$$rV = \frac{r + \lambda}{r + lambda + q_F}(-c - qm) + \tag{37}$$

$$\frac{q_F}{r+\lambda+q_F}\left[\left\{\int_{\underline{b}}^{b_m} E(x-m \mid R_m<x<x_{ml}) \times \frac{F(x_{ml})-F(R_m)}{1-F(R_m)} + E(x-w(x,b_l) \mid x>\hat{x_{ml}})\frac{1-F(x_{ml})}{1-F(R_m)} dH\right\}\right] +$$

$$\frac{q_F}{r+\lambda+q_F}\int_{b_m}^{\bar{b}} E(x-w(x,b_l) \mid x>R_l) dH$$

Now, consider matches involving a worker with $b < b_m$. In this case, rearranging (31) gives

$$R_m = rV + m \tag{38}$$

which implicitly defines R_m. We are interested in

$$\frac{\partial R_m}{\partial m} = r\frac{\partial V}{\partial R_m}\frac{\partial R}{\partial m} + r\frac{\partial V}{\partial m}\mid_{R_m} +1$$

Therefore

$$\frac{\partial R_m}{\partial m} = \frac{1+\frac{\partial rV}{\partial m}\mid_{R_m}}{1-r\frac{\partial V}{\partial R_m}} \tag{39}$$

First, consider $\frac{\partial V}{\partial R_m}$. An increase in R_m reduces the value of a vacancy because it means more initial matches are rejected (after incurring the search and training costs) but increases that value because the expected value of ongoing matches is now higher. By the envelope theorem, these effects will offset each other, implying $\frac{\partial V}{\partial R_m} = 0$ and,

$$\frac{R_m}{\partial m} = \frac{\partial rV}{\partial m}\mid_{R_m} +1$$

Now,

$$\frac{\partial rV}{\partial m}\mid_{R_m} = -\frac{r+\lambda}{r+\lambda+q_F}q+ \tag{40}$$

$$\frac{q_F}{r+\lambda+q_F}\left[\left\{\int_{\underline{b}}^{b_m} -1 * \frac{F(x_{ml})-F(R_m)}{1-F(R_m)} - \frac{\partial E(w(x,b_l) \mid x>x_{ml})}{\partial m}\frac{1-F(x_{ml})}{1-F(R_m)} dH\right\}\right] -$$

$$\frac{q_F}{r+\lambda+q_F}\int_{b_m}^{\bar{b}} \frac{\partial E(w(x,b_l) \mid \; x>x_{ml})}{\partial m} dH$$

To reduce this expression further, we need an expression for the bargained wage. To get this, note that we can express the surplus in a match with worker j and productivity draw x as,

$$S_j(x) = \frac{\eta - r(U_{uj}+V)}{r+\lambda} \tag{41}$$

The wage is set such that,

$$J_j(x) - V = (1-\beta) * S_j(x) \tag{42}$$

where, β is a parameter determining the relative bargaining power of the worker and firm. From this, we can write:

$$J_j(x) = (1-\beta)\frac{x - rU_{uj}}{r+\lambda} - (1-\beta)\frac{rV}{r+\lambda} + V \tag{43}$$

But from the basic firm Bellman equation we can also write:

$$J_j(x) = \frac{1}{r+\lambda}(x - w(x, b_j)) + \frac{\lambda}{r+\lambda}V \tag{44}$$

Setting (43) equal to (44) and solving, we get:

$$w(x, b_j) = rU_{uj} + \beta(x - rU_{uj} - rV) \tag{45}$$

Recalling that firms take worker reservation wages (rU_{uj}) as fixed, we then get:

$$\frac{\partial w_j}{\partial m} = -r\beta\frac{\partial V}{\partial m} \tag{46}$$

where we now write w_j instead of $w(x, b_j)$ since this derivative is not a function of x . We can now use this derivative in (40). To simplify notation, define,

$P_m = \int_{\underline{b}}^{b_m} \frac{F(x_{ml})-F(R_m)}{1-F(R_m)} dH$

as the probability a new match will end up being a minimum wage paying match and

$P_{Nm} = \int_{\underline{b}}^{b_m} \frac{1-F(x_{ml})}{1-F(R_m)} dH + \int_{b_m}^{\bar{b}} dH$

as the probability a new match will end up with a wage above the minimum wage. Then,

$$\frac{\partial rV}{\partial m} = -\frac{q(r+\lambda)}{r+\lambda+(1-\beta P_{Nm})q_F} - \frac{q_F}{r+\lambda+(1-\beta P_{Nm})q_F}P_m \tag{47}$$

which is negative. The first term on the right hand side corresponds to the increased hiring costs associated with paying workers the minimum wage during the probationary period. The second term corresponds to the fact that, with probability P_m a vacancy will ultimately filled with a minimum wage match and the profits of such matches are directly declining in the minimum wage. Thus, the sign of $\frac{\partial R_m}{\partial m}$ is uncertain. On one side, the increase in m implies the marginal x that just covers m plus the outside option of the firm is now higher. On the other side, once a match is formed (and the training cost paid), a rise in m implies a lower outside option for the firm. This will push the value for R_m down. In general, we would expect the direct effect to be larger than the second, indirect effect and, therefore, $\frac{R_m}{\partial m} > 0$.

Alternatively, for workers with $b > b_m$, rearranging (6) yields,

$$R_j = rV + w_j^* \tag{48}$$

In a partial equilibrium setting with w_j^* taken as fixed by firms, $\frac{\partial R_j}{\partial m}$ will be completely determined by $\frac{\partial rV}{\partial m}$. Since we have just seen that the latter derivative is negative, for matches with workers whose outside options are such that the wages paid in the match is above m, an increase in m leads to a decrease in layoffs. Whether the overall effect of an increase in m on layoffs is negative or positive then depends on the relative importance of the minimum wage versus non-minimum wage workers and is an empirical matter.

B2.2b Introducing Inflation

We turn next to allowing inflation in a partial equilibrium setting. To simplify the exposition and focus attention on minimum wage issues, we will assume that workers and firms are able to re-bargain at any point to maintain the real wage. Thus, the bargained wage $w(x, b_j)$ is a real wage, implying that inflation does not have a direct effect on the model outcome in the absence of minimum wages. However, minimum wages are re-set only sporadically by governments and so we will assume that the real value of the minimum wage is eroded by inflation over time. In particular, we will specify $m_t = m_0 e^{-\rho t}$, where m_t is the real minimum wage at time t, m_0 is the initial nominal minimum wage, and ρ is the inflation rate.

Given expected declines in the real minimum wage, the value of a match and the value of a vacancy to a firm both have a "capital gains" element to them reflecting expected changes in profit with future changes in the real minimum wage. Thus, dropping t subscripts for simplicity, we can re-write the firm's Bellman equations as:

$$rV = -c + q_F(J^e - V) - q * m + \dot{V} \tag{49}$$

$$rJ_j(x) = x - w(x, b_j) + \lambda(V - J_j(x)) + \dot{J}_j(x) \tag{50}$$

where, $\dot{V}$ and $\dot{J}_j(x)$ are the expected changes in the value of an unfilled and a filled vacancy, respectively. To re-iterate, these changes have to do only with anticipated changes in the real minimum wage due to inflation.

Consider a match with a worker whose reservation wage is above m. We can write,

$$R_j = rV + w_j^* - \dot{J}_j \tag{51}$$

and, therefore,

$$\frac{\partial R_j}{\partial m} = r\frac{\partial V}{\partial m} - \frac{\partial \dot{J}_j}{\partial m} \tag{52}$$

As before, to obtain an expression for $\frac{\partial V}{\partial m}$ we need to get an expression for expected profits from a filled match, J^e:

$$J^e = \frac{1}{r+\lambda}\int_{\underline{b}}^{b_m}\left[E(x \mid x > R_m) - m\frac{F(x_{ml}) - F(R_m)}{1 - F(R_m)} - E(w \mid x > x_{ml})\frac{1 - F(x_{ml})}{1 - F(R_m)}\right] dH \tag{53}$$

$$+\frac{1}{r+\lambda}\int_{b_m}^{\bar{b}}\left[E(x \mid x > R_l) - E(w \mid \ x > R_l)\right] dH + \frac{\delta}{r+\lambda}V + \frac{1}{r+\lambda}\dot{J}^e$$

Using this, we can write,

$$\frac{\partial rV}{\partial m} = -\frac{r+\lambda}{r+\lambda+q_F}q+ \tag{54}$$

$$\frac{q_F}{r+\lambda+q_F}\left[\left\{\int_{\underline{b}}^{b_m} -1 * \frac{F(x_{ml}) - F(R_m)}{1 - F(R_m)} - \frac{\partial E(w(x, b_l) \mid x > x_{ml})}{\partial m}\frac{1 - F(x_{ml})}{1 - F(R_m)}dH\right\}\right] -$$

$$\frac{q_F}{r+\lambda+q_F}\int_{b_m}^{\bar{b}}\frac{\partial E(w(x, b_l) \mid x > R_l)}{\partial m}dH + \frac{q_F}{r+\lambda+q_F}\frac{\partial \dot{j}^e}{\partial m} + \frac{r+\lambda}{r+\lambda+q_F}\frac{\partial \dot{V}}{\partial m}$$

As before, the derivative $\frac{\partial w_j}{\partial m}$ is crucial in evaluating this derivative. In this case it equals,

$$\frac{\partial w_j}{\partial m} = -\beta\frac{\partial rV}{\partial m} + \frac{\partial \dot{J}_j}{\partial m}. \tag{55}$$

Using this, we can write:

$$\frac{\partial rV}{\partial m} |_{R_m} = -\frac{r+\lambda}{r+\lambda+q_F} q \tag{56}$$

$$-\frac{q_F}{r+\lambda+q_F} P_m + \frac{q_F}{r+\lambda+q_F} \beta P_{Nm} \frac{\partial rV}{\partial m} - \frac{q_F}{r+\lambda+q_F} \left[\int_{\underline{b}}^{b_m} \frac{\partial \dot{J}_j}{\partial m} \frac{1-F(x_{ml})}{1-F(R_m)} dH + \int_{b_m}^{\bar{b}} \frac{\partial \dot{J}_j}{\partial m} dH \right]$$

$$+ \frac{q_F}{r+\lambda+q_F} \frac{\partial \dot{J}^e}{\partial m} + \frac{r+\lambda}{r+\lambda+q_F} \frac{\partial \dot{V}}{\partial m}$$

To evaluate this further, we need to get an expression for $\dot{J}_j$. Given that future values of a match vary only because of inflation and inflation is assumed to have an impact only through changes in the real minimum wage, we have that,

$$\dot{J}_j = \frac{\partial J_j}{\partial m} \frac{\partial m}{\partial t} = \frac{\partial J_j}{\partial m} (-\rho m_0 e^{-\rho t}) \tag{57}$$

Using the definition of J_j from the firm's Bellman equation, this implies that for a match that pays a wage above m:

$$\dot{J}_j = \frac{1}{r+\lambda} (-\rho m_0 e^{-\rho t}) \left[-\frac{\partial w_j}{\partial m} + \lambda \frac{\partial V}{\partial m} + \frac{\partial \dot{J}^e}{\partial m} \right] \tag{58}$$

$$= \frac{1}{r+\lambda} (-\rho m_0 e^{-\rho t}) \left[(\lambda + \beta r) \frac{\partial V}{\partial m} \right]$$

Essentially, as inflation drives down the real minimum wage, and as a result drives up V, the negotiated wage will decline and profits rise.

The derivative $\frac{\partial \dot{J}_j}{\partial m}$ in (56) involves taking the derivative of (58) with respect to m_0:

$$\frac{\partial \dot{J}_j}{\partial m_0} = \frac{-\rho e^{-\rho t}}{r+\lambda} (\lambda + \beta r) \frac{\partial V}{\partial m} - \frac{\rho m_0 e^{-\rho t}}{r+\lambda} (\lambda + \beta r) \frac{\partial^2 V}{\partial m^2} \tag{59}$$

We also need an expression for $\frac{\partial \dot{J}^e}{\partial m}$. To get this, note first that:

$$\dot{J}^e = \frac{-\rho m_0 e^{-\rho t}}{r+\lambda} \left[\int_{\underline{b}}^{b_m} \left(-1 + \lambda \frac{\partial V}{\partial m} + \frac{\partial \dot{J}_j}{\partial m} \right) \frac{F(x_{ml}) - F(R_m)}{1-F(R_m)} dH + P_{Nm} (\lambda + \beta r) \frac{\partial V}{\partial m} \right] \tag{60}$$

To evaluate this, we need an expression for $\frac{\partial \dot{J}_j}{\partial m}$ for jobs that pay the minimum wage:

$$\frac{\partial \dot{J}_j}{\partial m_0} = \frac{\rho e^{-\rho t}}{r+\lambda+\rho e^{-\rho t}} * \left[1 - \lambda \frac{\partial V}{\partial m} - m_0 \lambda \frac{\partial^2 V}{\partial m^2} - \frac{\partial^2 \dot{J}_j}{\partial m_0^2} \right] \tag{61}$$

This is a differential equation that yields the solution:

$$\frac{\partial \dot{J}_j}{\partial m_0} = \frac{\rho e^{-\rho t}}{r+\lambda+\rho e^{-\rho t}} (1 - \lambda \frac{\partial V}{\partial m}) - \frac{\rho e^{-\rho t}}{r+\lambda+2\rho e^{-\rho t}} \lambda \frac{\partial^2 V}{\partial m^2} m + \kappa_1 m^{-\frac{r+\lambda+\rho e^{-\rho t}}{\rho e^{-\rho t}}} \tag{62}$$

where, κ_1 is an arbitrary constant. Note that if there is no inflation (i.e., $\rho = 0$) then $\frac{\partial \dot{J}_j}{\partial m_0}$ should equal zero. To insure this is the true, we need to set $\kappa_1 = 0$.

Plugging (62) into (60) yields an expression for $\frac{\partial j_j}{\partial m}$ in terms of $\frac{\partial V}{\partial m}, and \frac{\partial^2 V}{\partial m^2} m$. Taking the derivative of that expression with respect to m will then generate an expression for $\frac{\partial j^e}{\partial m}$ in terms of $\frac{\partial V}{\partial m}, \frac{\partial^2 V}{\partial m^2}, \frac{\partial^2 V}{\partial m^2} m, and \frac{\partial^3 V}{\partial m^3} m$. Substituting this expression and (59) into (56) yields an expression:

$$r\frac{\partial V}{\partial m} = -\frac{(r+\lambda)\phi}{r+\lambda+q_F} - aP_m + abP_m - abP_m c \tag{63}$$

$$+\left[aP_{Nm}\beta r - abP_{Nm}(\lambda+\beta r) - ab(\lambda + P_{Nm}\beta r) + abP_m c\lambda - \frac{(r+\lambda)\rho e^{-\rho t}}{r+\lambda+q_F}\right]\frac{\partial V}{\partial m}$$

$$+\left[(-aP_{Nm}b(\lambda+\beta r) + abP_m d\lambda - ab(\lambda + P_{Nm}\beta r) + abc\lambda P_m + abd\lambda P_m - \frac{(r+\lambda)\rho e^{-\rho t}}{r+\lambda+q_F}\right]\frac{\partial^2 V}{\partial m^2}$$

$$+abd\delta P_m \frac{\partial^3 V}{\partial m^3}$$

where, $a = \frac{q_F}{r+\lambda+q_F}, b = \frac{\rho e^{-\rho t}}{r+\lambda}, c = \frac{\rho e^{-\rho t}}{r+\lambda+\rho e^{-\rho t}}, and, d = \frac{\rho e^{-\rho t}}{r+\lambda+2\rho e^{-\rho t}}$.
Summarizing further, we can write (63) as:

$$A\frac{\partial V}{\partial m} + B + C\frac{\partial^2 V}{\partial m^2} m + D\frac{\partial^3 V}{\partial m^3} m^2 \tag{64}$$

where A, B, C, and D are all functions of parameters. This is a second order differential equation for $\frac{\partial V}{\partial m}$ in m, and its solution is:

$$\frac{\partial V}{\partial m} = \frac{B}{A} + (\kappa_2 + \kappa_3) m^{\frac{-(4AD+C^2-2CD+D^2)^{0.5}-C+D}{2D}} \tag{65}$$

We can write B in this expression as,

$$B = -\frac{(r+\lambda)q}{r+\lambda+q_F} - aP_m + aP_m \frac{\rho e^{-\rho t}}{r+\lambda+\rho e^{-\rho t}} \tag{66}$$

and

$$A = (r - aP_{Nm}\beta r) + abP_{Nm}(2\beta r + \lambda) + ab\lambda(1 - cP_m) + \frac{(r+\lambda)\rho e^{-\rho t}}{r+\lambda+q_F} \tag{67}$$

If there is no inflation (i.e., if $\rho = 0$) then B/A reduces to (47), the expression for $\frac{\partial V}{\partial m}$ in the absence of inflation. In particular, the numerators in the first two terms in (66) are the same as the numerators in the terms on the right hand side of (47). The third term in (66) is positive if $\rho > 0$. Essentially, the second term in (66) corresponds to the negative effect of an increase in m on V because of the reduction in profits from future potential minimum wage matches. This effect is larger the greater the probability of such a match and is discounted according to the discount rate and λ (the probability any such match would be terminated exogenously). The third term corresponds to a reduction in this discounted effect to the extent there is inflation. At the same time, the numerators in the first two terms in A correspond to the denominator in (47). If $\rho > 0$ the denominator becomes larger and more positive. Thus, with $\rho > 0$ the ratio B/A in (65) has a numerator that is closer to zero and a denominator that is more positive than (47), implying that with inflation, $\frac{\partial V}{\partial m}$ is larger (i.e., less negative) than without inflation.

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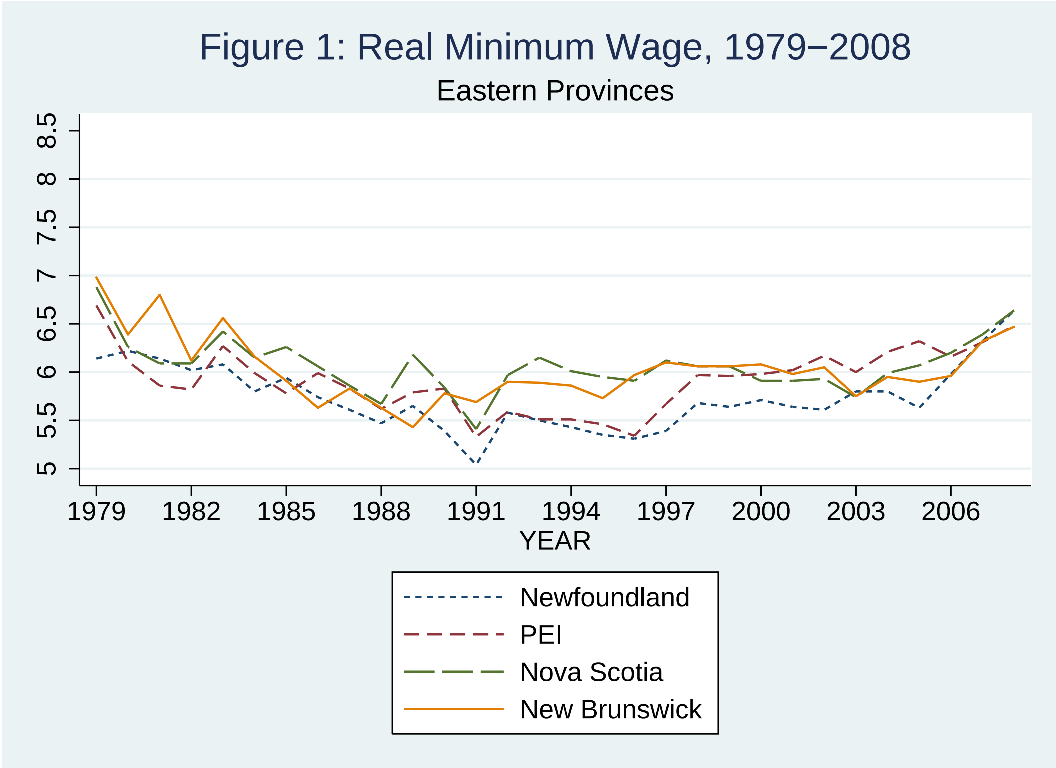
Figure 1: Real Minimum Wage, 1979−2008
Eastern Provinces
5
5.5
6
6.5
7
7.5
8
8.5
1979
1982
1985
1988
1991
1994
1997
2000
2003
2006
YEAR
Newfoundland
PEI
Nova Scotia
New Brunswick

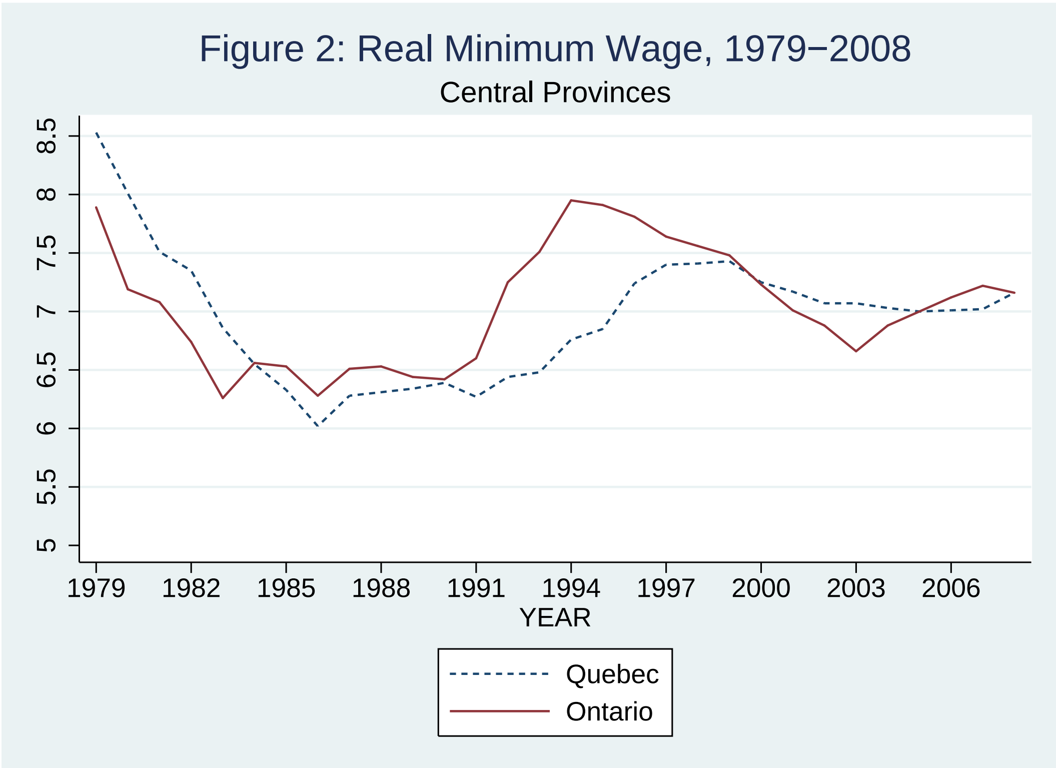
Figure 2: Real Minimum Wage, 1979−2008
Central Provinces
5
5.5
6
6.5
7
7.5
8
8.5
1979
1982
1985
1988
1991
1994
1997
2000
2003
2006
YEAR
Quebec
Ontario

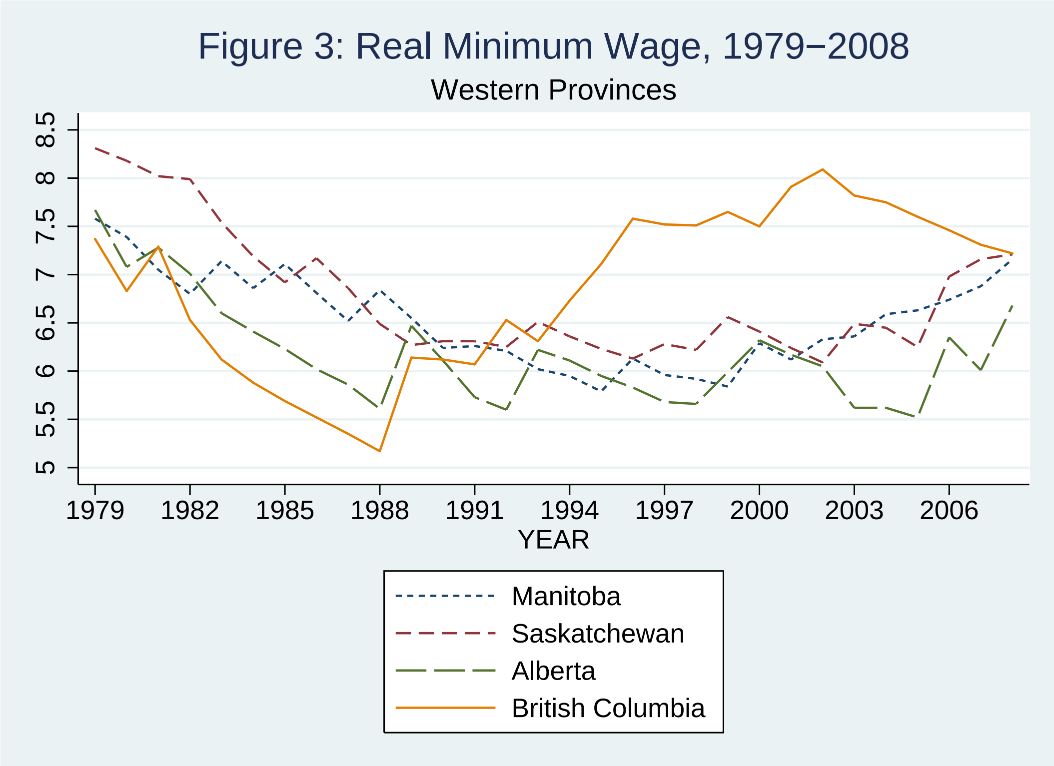
Figure 3: Real Minimum Wage, 1979−2008
Western Provinces
5
5.5
6
6.5
7
7.5
8
8.5
1979
1982
1985
1988
1991
1994
1997
2000
2003
2006
YEAR
Manitoba
Saskatchewan
Alberta
British Columbia

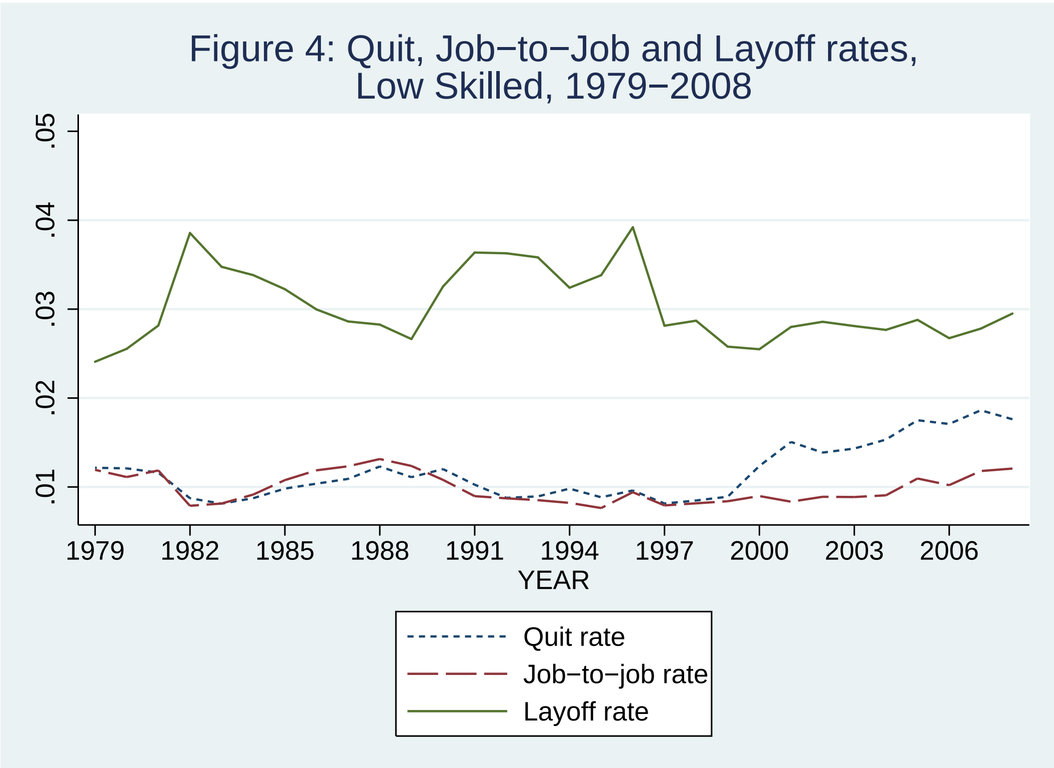
Figure 4: Quit, Job−to−Job and Layoff rates,
Low Skilled, 1979−2008
.01
.02
.03
.04
.05
1979
1982
1985
1988
1991
1994
1997
2000
2003
2006
YEAR
Quit rate
Job−to−job rate
Layoff rate

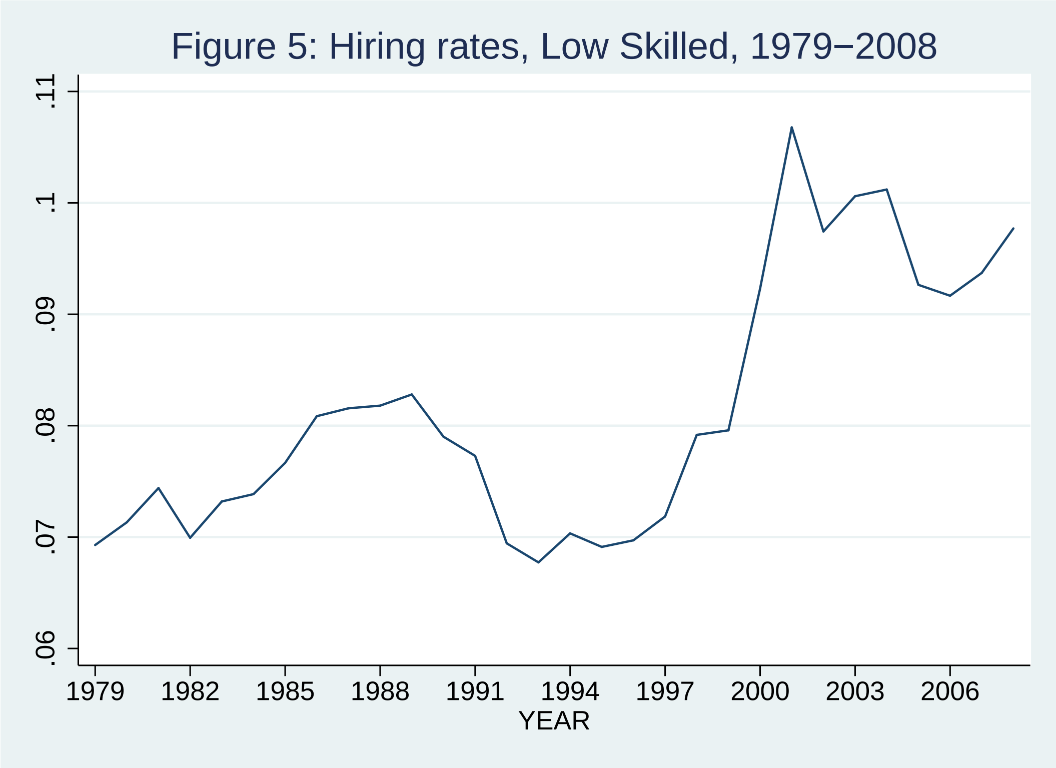
Figure 5: Hiring rates, Low Skilled, 1979−2008
.06
.07
.08
.09
.1
.11
1979
1982
1985
1988
1991
1994
1997
2000
2003
2006
YEAR

Table 1: Separation Rate, Low Skilled

	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
Males and Females				
No Lags				
lrmin	-.0009	-.0351	-.0434	-.0075
	(.0053)	(.0114)***	(.0134)***	(.0107)
R-squared	.65	.65	.62	.51
With 1 Year Lag				
lrmin	-.0028	-.0318	-.0419	.0010
	(.0081)	(.0173)*	(.0202)**	(.0180)
lrminlag12m	.0024	-.0031	.0002	-.0107
	(.0081)	(.0172)	(.0202)	(.0180)
R-squared	.66	.66	.63	.52
Males				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	-.0011	-.0380	-.0436	-.0107
	(.0057)	(.0133)***	(.0156)***	(.0138)
R-squared	.65	.62	.57	.51
With 1 Year Lag				
lrmin	-.0066	-.0396	-.0444	-.0151
	(.0089)	(.0209)*	(.0245)*	(.0235)
lrminlag12m	.0063	.0013	.0022	.0028
	(.0089)	(.0208)	(.0244)	(.0234)
R-squared	.66	.63	.57	.52
Females				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	-.0009	-.0318	-.0420	-.0061
	(.0060)	(.0127)***	(.0153)***	(.0119)
R-squared	.57	.56	.54	.34
With 1 Year Lag				
lrmin	.0009	-.0282	-.0471	.0117
	(.0093)	(.0198)	(.0240)**	(.0201)
lrminlag12m	-.0018	-.0018	.0081	-.0200
	(.0093)	(.0198)	(.0241)	(.0202)
R-squared	.57	.57	.54	.35

Notes. Dependent variable: proportion of workers on a job in month t who separate from that job in month t+1. lrmin is the log of the real minimum wage. All regressions are estimated using FGLS (AR(3) model), and are weighted by the inverse of the number in the at-risk group. The number of observations is 3,472 in specifications without a lag and 3,392 in specifications with a lag. All regressions include a full set of time and province dummies and a dummy equal to one if there was a minimum wage change in the month. Standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table 2: Separation Rate by Age Group, Low Skilled

	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
15 to 19 Years of Age				
No Lags				
lrmin	-.0379	-.0404	-.0513	.0028
	(.0133)***	(.0166)***	(.0201)***	(.0210)
R-squared	.50	.46	.41	.21
With 1 Year Lag				
lrmin	-.0032	-.0158	-.0307	.0225
	(.0218)	(.0274)	(.0333)	(.0362)
lrminlag12m	-.0357	-.0227	-.0149	-.0250
	(.0220)	(.0275)	(.0334)	(.0364)
R-squared	.49	.46	.40	.21
20 to 24 Years of Age				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	.0010	-.0175	-.0273	.0058
	(.0078)	(.0128)	(.0161)*	(.0150)
R-squared	.53	.49	.44	.27
With 1 Year Lag				
lrmin	-.0058	-.0312	-.0325	-.0194
	(.0127)	(.0213)	(.0269)	(.0259)
lrminlag12m	.0075	.0170	.0053	.0318
	(.0128)	(.0213)	(.0269)	(.0259)
R-squared	.54	.49	.44	.27
25 to 59 Years of Age				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	.0008	-.0347	-.0404	-.0123
	(.0055)	(.0139)***	(.0161)**	.0129
R-squared	.61	.62	.58	.52
With 1 Year Lag				
lrmin	-.0025	-.0329	-.0491	.0051
	(.0082)	(.0208)	(.0243)**	(.0215)
lrminlag12m	.0040	-.0029	.0111	-.0224
	(.0082)	(.0207)	(.0243)	(.0215)
R-squared	.62	.63	.59	.52

Notes. Dependent variable: proportion of workers on a job in month t who separate from that job in month t+1. lrmin is the log of the real minimum wage. All regressions are estimated using FGLS (AR(3) model), and are weighted by the inverse of the number in the at-risk group. The number of observations is 3,472 in specifications without a lag and 3,392 in specifications with a lag. All regressions include a full set of time and province dummies and a dummy equal to one if there was a minimum wage change in the month. Standard errors in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table 3: Quit, Job-to-Job, and Layoff rates, Low Skilled

Quits				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	.0015	-.0039	-.0031	-.0019
	(.0011)	(.0026)	(.0035)	(.0033)
R-squared	.64	.56	.53	.30
With 1 Year Lag				
lrmin	.0042	.0042	.0024	.0088
	(.0018)**	(.0043)	(.0058)	(.0057)
lrminlag12m	-.0035	-.0094	-.0059	-.0131
	(.0018)*	(.0043)**	(.0058)	(.0057)**
R-squared	.65	.56	.53	.30
Layoffs				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	-.0054	-.0253	-.0326	-.0055
	(.0050)	(.0017)**	(.0133)***	(.0101)
R-squared	.64	.67	.64	.58
With 1 Year Lag				
lrmin	-.0085	-.0268	-.0303	-.0100
	(.0075)	(.0172)	(.0196)	(.0169)
lrminlag12m	.0036	.0013	-.0032	.0043
	(.0075)	(.0171)	(.0195)	(.0169)
R-squared	.65	.67	.65	.58
Job to Job Transitions				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	-.0024	-.0142	-.0172	-.0062
	(.0013)*	(.0030)***	(.0040)***	(.0031)**
R-squared	.56	.48	.45	.22
With 1 Year Lag				
lrmin	.0003	-.0072	-.0106	.0021
	(.0020)	(.0049)	(.0066)	(.0054)
lrminlag12m	-.0036	-.0085	-.0084	-.0098
	(.0020)*	(.0049)*	(.0066)	(.0054)*
R-squared	.56	.47	.44	.22

Notes. Dependent variable: proportion of workers on a job in month t who leave the job by each route. lrmin is the log of the real minimum wage. All regressions are estimated using FGLS (AR(3) model), and are weighted by the inverse of the number in the at-risk group. The number of observations is 3,472 in specifications without a lag and 3,392 in specifications with a lag. All regressions include a full set of time and province dummies and a dummy equal to one if there was a minimum wage change in the month. Standard errors in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table 4: Hiring Rate, Out of LF to unemployment transition, Unemployment to out out of LF transition, and Hours of Work, Low Skilled

Hiring				
No Lags				
	Both Genders	Males	Females	Teenagers Both Genders
lrmin	-.0270	-.0349	-.0145	-.0846
	(.0077)***	(.0144)***	(.0063)**	(.0217)***
R-squared	.57	.58	.53	.40
With 1 Year Lag				
lrmin	-.0140	-.0101	-.0078	-.0371
	(.0126)	(.0227)	(.0102)	(.0334)
lrminlag12m	-.0163	-.0333	-.0091	-.0672
	(.0127)	(.0227)	(.0103)	(.0336)**
R-squared	.56	.57	.53	.40
Out of LF to Unemployment transition			**Unemployment to out of LF transition**	
No Lags			*No Lags*	
	Both Genders			Both Genders
lrmin	-.0194		lrmin	.0250
	(.0064)***			(.0149)
R-squared	.48		R-squared	.39
With 1 Year Lag				
lrmin	-.0163		lrmin	-.0120
	(.0096)*			(.0225)
lrminlag12m	-.0023		lrminlag12m	.0492
	(.0096)			(.0225)**
R-squared	.47		R-squared	.38
Hours of Work				
No Lags				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	.0003	.0019	.0015	.0039
	(.0026)	(.0043)	(.0058)	(.0055)
R-squared	.83	.74	.70	.66

Notes. Dependent variables: proportion of non-employed in month t who find a job in t+1. Proportion out of the labour force in month t who are unemployed in t+1, proportion of unemployed in month t that are out of the labour force in t+1, and the change in average weekly hours. lrmin is the log of the real minimum wage. All regressions are estimated using FGLS (AR(3) model), and are weighted by the inverse of the number in the at-risk group. The number of observations is 3,472 in specifications without a lag and 3,392 in specifications with a lag. All regressions include a full set of time and province dummies and a dummy equal to one if there was a minimum wage change in the month. Standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table 5: Robustness Checks, <6 months tenure

Separation Rate							
	Base	Min.Wage Ratio	IV Nominal Minimum	Drop Post-Announce	Males Age 25-54 with BA	Drop Recessions	IV Political
lrmin	-.043	-.041	-.052	-.047	-.019		-.13
	(.013)***	(.0099)***	(.024)**	(.016)***	(.020)***		(.045)***
Layoff Rate							
	Base	Min.Wage Ratio	IV Nominal Minimum	Drop Post-Announce	Males Age 25-54 with BA	Drop Recessions	IV Political
lrmin	-.033	-.046	-.041	-.038	-.0001	-.036	-.085
	(.013)***	(.0096)***	(.019)**	(.015)***	(.015)***	(.015)**	(.036)**

Notes. Dependent variables: proportion of workers on a job in month t who separate from that job in month t+1 and proportion who are laid off by month t+1. lrmin represents the log of the real minimum wage in all columns except column 2 where it corresponds to the ratio of (the minimum wage times 40) to the median weekly wage for males with a high school or less education. The median wage variable varies only at the annual level. Column 4 results are based on a sample dropping all months between an announcement of a minimum wage increase and the actual increase. Column 6 is run using only non-recessionary months (1979m1 to 1981m12, 1986m9 to 1990m11, and 1997m1 to 2007m12). The instruments in the last column consist of a dummy if the governing party is Left wing, a dummy for whether the governing party is Right wing, and the averages of each of these variables for all other provinces in the region. All regressions are estimated using FGLS (AR(3) model) except for the IV specifications where standard errors clustered at the provincial level are reported, and are weighted by the inverse of the number in the at-risk group. All regressions include a full set of time and province dummies and a dummy equal to one if there was a minimum wage change in the month. Standard errors in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table 6: Separation, Quit, Job-to-Job, and Layoff Rates with Inflation Interaction, Low Skilled

Separation				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	-.0158	-.0652	-.0755	-.0240
	(.0067)***	(.0144)***	(.0169)***	(.0141)*
inf	-.0073	-.0150	-.0164	-.0079
	(.0021)***	(.0045)***	(.0052)***	(.0045)*
lrmin*inf	.0037	.0070	.0074	.0043
	(.0011)***	(.0024)***	(.0027)***	(.0024)*
R-squared	.67	.67	.63	.52
Quits				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	.0009	-.0013	.0025	-.0046
	(.0014)	(.0034)	(.0046)	(.0045)
inf	-.0002	.0010	.0024	-.0013
	(.0005)	(.0011)	(.0014)*	(.0014)
lrmin*inf	.0002	-.0003	-.0010	.0009
	(.0002)	(.0006)	(.0007)	(.0008)
R-squared	.65	.56	.53	.0008
Job-to-Job Transitions				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	-.0061	-.0205	-.0244	-.0108
	(.0016)***	(.0039)***	(.0053)***	(.0042)***
inf	-.0017	-.0031	-.0035	-.0024
	(.0005)***	(.0012)***	(.0016)***	(.0014)*
lrmin*inf	.0010	.0017	.0019	.0014
	(.0003)***	(.0006)***	(.0009)**	(.0007)**
R-squared	.57	.47	.45	.22
Layoffs				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
lrmin	-0.0195	-.0573	-.0711	-.0188
	(.0062)***	(.0147)***	(.0167)***	(.0134)
inf	-.0067	-.0150	-.0185	-.0060
	(.0020)***	(.0045)***	(.0051)***	(.0043)
lrmin*inf	.0032	.0070	.0084	.0030
	(.0011)***	(.0024)***	(.0027)***	(.0023)
R-squared	.66	.68	.66	.58

Notes. Dependent variable: proportion of workers on a job in month t who leave the job by each route. lrmin is the log of the real minimum wage. All regressions are estimated using FGLS (AR(3) model), and are weighted by the inverse of the number in the at-risk group. The number of observations is 3,472 in specifications without a lag and 3,392 in specifications with a lag. All regressions include a full set of time and province dummies and a dummy equal to one if there was a minimum wage change in the month. Standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01.

Table A.1: Mean Separation, Quit, Job-to-Job, Layoff Rates and Hours Growth, Low Skilled

	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
Separation				
Overall	.0562	.1321	.1601	.0894
Males	.0581	.1415	.1690	.0964
Females	.0541	.1210	.1486	.0819
Teenagers	.1423	.1707	.2015	.1108
Young Adults	.0858	.1325	.1598	.0908
Older Adults	.0460	.1229	.1487	.0849
Quit				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
Overall	.0117	.0257	.0304	.0185
Males	.0096	.0226	.0275	.0150
Females	.0143	.0293	.0342	.0225
Teenagers	.0406	.0474	.0545	.0337
Young Adults	.0221	.0315	.0366	.0236
Older Adults	.0083	.0184	.0216	.0138
Direct Job to Job Transitions				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
Overall	.0099	.0264	.0332	.0160
Males	.0111	.0302	.0377	.0180
Females	.0084	.0219	.0276	.0138
Teenagers	.0344	.0401	.0477	.0253
Young Adults	.0194	.0308	.0380	.0199
Older Adults	.0068	.0216	.0275	.0128
Layoff				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
Overall	.0303	.0742	.0894	.0506
Males	.0337	.0834	.0975	.0598
Females	.0264	.0632	.0789	.0407
Teenagers	.0594	.0747	.0894	.0460
Young Adults	.0396	.0648	.0788	.0432
Older Adults	.0270	.0775	.0931	.0542
Hours Growth				
	All Tenure	< 1 Year	< 6 Months	6 to 11 Months
Overall	-.0030	-.0030	-.0017	-.0050
Males	-.0045	-.0051	-.0038	-.0073
Females	-.0008	.0010	.0017	-.0020
Teenagers	-.0064	-.0056	-.0054	-.0063
Young Adults	-.0036	-.0032	-.0013	-.0060
Older Adults	-.0028	-.0025	-.0011	-.0046

Table A.2: Mean Hiring, UN and NU Rates, Low Skilled

Hiring (Conditioning on Being Initially Out of Work)	
Overall	.0814
Males	.1282
Females	.0578
Teenagers	.1611
Young Adults	.1423
Older Adults	.0682
UN	
Overall	.1605
NU	
Overall	.0656

Table A.3: Proportions in the Various Tenure Categories for each Age Group, Low Skilled

Age Groups			
	< 6 Months	6 to 11 Months	≥ 1 Year
Teenagers	.4399	.2298	.3303
Young Adults	.2561	.1707	.5731
Older Adults	.1048	.0740	.8212