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## The cyclical behavior of equilibrium unemployment and vacancies revisited

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**WORKING PAPER SERIES**

**NO 853 / JANUARY 2008**

**THE CYCLICAL  
BEHAVIOR OF  
EQUILIBRIUM  
UNEMPLOYMENT  
AND VACANCIES  
REVISITED**

by Marcus Hagedorn  
and Iourii Manovskii



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# THE CYCLICAL BEHAVIOR OF EQUILIBRIUM UNEMPLOYMENT AND VACANCIES REVISITED<sup>1</sup>

by Marcus Hagedorn<sup>2</sup>  
and Iouri Manovskii<sup>3</sup>



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## **Abstract**

Recently, a number of authors have argued that the standard search model cannot generate the observed business-cycle-frequency fluctuations in unemployment and job vacancies, given shocks of a plausible magnitude. We use data on the cost of vacancy creation and cyclicalities of wages to identify the two key parameters of the model - the value of non-market activity and the bargaining weights. Our calibration implies that the model is, in fact, consistent with the data.

JEL Classification: E24, E32, J41, J63, J64

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## Non-Technical Summary

The unemployment rate is one of the key macroeconomic variables. In most industrialized countries, the unemployment series is strongly countercyclical: it rises sharply in a recession and declines in a boom. Similarly, the number of vacancies, or jobs looking for workers, also varies substantially over the business cycle. It rises sharply in a boom and declines in recessions.

The leading theory of equilibrium unemployment and vacancies is the Mortensen-Pissarides model, which explains the co-existence of unemployment and vacancies through frictions in matching workers and potential employers. One of the appealing features of this model is that generates the predictions that have the right direction: unemployment in the model indeed goes up in recessions and down in boom, while job vacancies move in the opposite direction.

However, it has been argued recently that the model cannot match the size of these variations. This was a very important set back for the profession. The reason is that this is the workhorse model that is used by academic and government economists to evaluate various economic policies and to study the problem of unemployment caused by the business cycles. A huge literature is developing that attempts to modify the basic model in order to make it consistent with the data on unemployment and vacancies fluctuations.

In this paper we argue that the ability of the model to generate large fluctuations in unemployment and vacancies crucially depends on the way the key parameters of the model are chosen. We use data on the cost of posting vacancies and the cyclicalities of wages to identify these key parameters. Our strategy implies that the model is, in fact, consistent with labour market data.

The explanation for our result is that our model generates large percentage fluctuations in profits. In a boom, there is a large percentage increase in profits and firms find it profitable to create many jobs. In a recession, profits drop and job creation is abnormally low. The reason for these large percentage fluctuations is that we find profits to be small (relative to GNP) and that wages do not respond one-for-one to changes in productivity over the business cycle.

# 1 Introduction

The Mortensen-Pissarides (MP) search and matching model (Mortensen and Pissarides (1994), Pissarides (1985, 2000)) has become the standard theory of equilibrium unemployment. It provides an appealing description of the labor market and has been found relevant in quantitative work. For example, Merz (1995) and Andolfatto (1996) have shown that the performance of the real business cycle model can be improved significantly when the MP model is embedded into it. However, Andolfatto (1996) and Shimer (2005) have argued that the standard calibration of the model fails to account for the cyclical properties of its two central variables - unemployment and vacancies. These variables are much more volatile in U.S. data than in the MP model.

The literature has responded by suggesting that the wage setting mechanism in the MP model has to be altered to break the close link between wages and productivity. Farmer (2004) and Shimer (2004) suggest that some rigidity in wage formation may be necessary. In Hall (2005a) and Gertler and Trigari (2005) a form of social wage norm renders wages not responsive to productivity changes. Hall and Milgrom (2006) modify the bargaining game to limit the influence of labor market conditions on wages. Kennan (2005) and Menzio (2004) endogenize wage rigidity by modeling asymmetric information about productivity.<sup>1</sup>

We take a different route in this paper. We suggest that the problem lies not in the model itself, but in the way the model is typically calibrated. We propose a new calibration strategy for the two central parameters of the MP model - the worker's value of non-market activity and the worker's bargaining power. We measure the costs of posting vacancies and the cyclicity of wages in the data to pin down these two key parameters. Instead, the usual strategy is to identify non-market activity with receiving unemployment benefits. The bargaining weight is then picked in a way that guarantees the efficiency of the model

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<sup>1</sup>Many other authors have explored whether a search model is consistent with business cycle facts, including Lilien (1982), Abraham and Katz (1986), Blanchard and Diamond (1989), Merz (1995), Ramey and Watson (1997), Cole and Rogerson (1999), Den Haan, Ramey, and Watson (2000), Gomes, Greenwood, and Rebelo (2001), Fujita (2004), Pries (2004), Brügemann and Moscarini (2005), Costain and Reiter (2005). See Hornstein, Krusell, and Violante (2005b) and Yashiv (2006) for recent surveys of the literature.



(i.e., to satisfy the Hosios (1990) condition). This choice of the bargaining power implies that wages in the model follow productivity closely over the business cycle. The problem is then clear. An increase in productivity is mostly absorbed by an increase in wages and profits remain little changed. Therefore, firms' incentives to post vacancies do not increase much either.

Our calibration strategy does not impose the assumption that the return to non-market activity is identical to receiving unemployment benefits. Indeed, in a model without search frictions, for example, in a standard real business cycle model, market and non-market productivities are equalized: workers are indifferent between working one more hour at home or in the market in Benhabib, Rogerson, and Wright (1991) and Greenwood and Hercowitz (1991) and value equally market and non-market activities in Hansen (1985) and Rogerson (1988). Thus, it seems arbitrary to assume a large gap between the value of market and non-market activities. It also appears more natural to use data to identify bargaining weights rather than the efficiency condition. In fact, we show that our calibration of the model that incorporates U.S. tax rates identifies the value for the bargaining weight that is much closer to the efficient benchmark than the value used in the standard calibration.

In the MP model firms incur costs of posting a vacancy and recover these costs by paying workers less than their marginal product. This gives rise to the period-by-period accounting profits. Free entry ensures that expected economic profits from posting are zero. For firms' vacancy posting decisions to respond strongly to changes in productivity the rewards that firms get for posting (i.e., the expected profits) have to change sharply with changes in productivity. For small changes in productivity to result in large percentage changes in profits, profits have to be small and strongly procyclical.

We measure the costs of posting vacancies in the data and find that they are small, implying small accounting profits in the calibrated model. The MP model is consistent with small profits either when the value of non-market activity of the worker is high, or when it is low but the worker has a large bargaining weight. The fact that wages are only moderately procyclical in the data uniquely pins down the worker's bargaining weight at a relatively low value, implying a value of non-market activity in the model that is considerably higher

than the typical replacement ratio of unemployment insurance. Our calibration of the model implies that it is consistent with the cyclical volatility of unemployment and vacancies.

The paper is organized as follows. A discrete time stochastic version of the Pissarides (1985, 2000) search and matching model is laid out in Section 2. In Section 3 we develop our calibration strategy, perform a quantitative analysis, and evaluate its robustness. We show that the model accounts very well for the volatility of unemployment and vacancies when we pin down the value of non-market activity and worker's bargaining power by matching the data on the costs of vacancy creation and on the cyclicalities of wages. In Section 4 we discuss several implications of our calibration approach. Section 5 concludes.

## 2 The Model

We consider a stochastic discrete time version of the Pissarides (1985, 2000) search and matching model with aggregate uncertainty. The exposition of the model follows Shimer (2005). The only differences are that from the outset we write down the discrete time model that we will use in quantitative analysis, we add capital to the model, and we allow the cost of posting vacancies to vary over the business cycle.

### 2.1 Workers and Firms

There is a measure one of infinitely lived workers and a continuum of infinitely lived firms. Workers maximize their expected lifetime utility:

$$\mathbb{E} \sum_{t=0}^{\infty} \delta^t y_t, \quad (1)$$

where  $y_t$  represents income in period  $t$  and  $\delta \in (0, 1)$  is workers' and firms' common discount factor. Firms have a constant returns to scale production technology that, for now, uses labor as the only input (we will add capital to the model in Section 2.4). Output of each unit of labor is denoted by  $p_t$ . Labor productivity  $p_t$  follows a first order Markov process in discrete time, according to some distribution  $G(p', p) = Pr(p_{t+1} \leq p' \mid p_t = p)$ .

There is free entry of firms. Firms attract unemployed workers by posting a vacancy at the flow cost  $c_p$ . Throughout the paper the notation  $X_p$  indicates that a variable  $X$  is a

function of the aggregate productivity level  $p$ . Once matched, workers and firms separate exogenously with probability  $s$  per period (see Hall (2005b) for the evidence that  $s$  is constant over the business cycle). Employed workers are paid a wage  $w_p$ , and firms make accounting profits of  $p - w_p$  per worker each period in which they operate. Unemployed workers get flow utility  $z$  from leisure/non-market activity.<sup>2</sup> Workers and firms split the surplus from a match according to the generalized Nash bargaining solution. The bargaining power of workers is  $\beta \in (0, 1)$ .

## 2.2 Matching

Let  $u_t$  denote the unemployment rate (or the number of unemployed people) and  $n_t = 1 - u_t$  the employment rate. Let  $v_t$  be the number of vacancies posted in period  $t$ . We refer to  $\theta_t = v_t/u_t$  as the market tightness at time  $t$ .

The number of new matches (starting to produce output at  $t + 1$ ) is given by a constant returns to scale matching function  $m(u_t, v_t)$ . Employment evolves according to the following law of motion:

$$n_{t+1} = (1 - s)n_t + m(u_t, v_t). \quad (2)$$

The probability for an unemployed worker to be matched with a vacancy next period equals  $f(\theta_t) = m(u_t, v_t)/u_t = m(1, \theta_t)$ . The probability for a vacancy to be filled next period equals  $q(\theta_t) = m(u_t, v_t)/v_t = m(1/\theta_t, 1) = f(\theta_t)/\theta_t$ . We restrict  $m(u_t, v_t) \leq \min(u_t, v_t)$ .

## 2.3 Equilibrium

Denote the firm's value of a job (a filled vacancy) by  $J$ , the firm's value of an unfilled vacancy by  $V$ , the worker's value of having a job by  $W$ , and the worker's value of being unemployed by  $U$ . Let  $E_p X_{p'}$  denote next period's expected value of an arbitrary variable

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<sup>2</sup> Without search, indivisibility of labor implies  $p = z$  in equilibrium (Hansen (1985), Rogerson (1988)). Consider a family of measure one. The family decides what fraction of its members,  $L$ , should work in the market, given that each worker can produce  $z$  at home, to  $\max_L \{Lp + (1 - L)z\}$ , where  $p = F_L(L, K)$  denotes the marginal product of labor. Assuming an interior solution, the optimal choice of  $L$  implies  $p = z$ .

$X$ , conditional on the current state  $p$ . With this notation, the following Bellman equations describe the model:<sup>3</sup>

$$J_p = p - w_p + \delta(1 - s)E_p J_{p'} \quad (3)$$

$$V_p = -c_p + \delta q(\theta_p)E_p J_{p'} \quad (4)$$

$$U_p = z + \delta\{f(\theta_p)E_p W_{p'} + (1 - f(\theta_p))E_p U_{p'}\} \quad (5)$$

$$W_p = w_p + \delta\{(1 - s)E_p W_{p'} + sE_p U_{p'}\}. \quad (6)$$

The interpretation is straightforward. Operating firms earn profits  $p - w_p$  and the matches are exogenously destroyed with probability  $s$ . A vacancy costs  $c_p$  and is matched with a worker (becomes productive next period) with probability  $q(\theta_p)$ . An unemployed worker derives utility  $z$  and finds a job next period with probability  $f(\theta_p)$ . An employed worker earns wage  $w_p$  but may lose her job with probability  $s$  and become unemployed next period.

Nash bargaining implies that a worker and a firm split the surplus  $S_p = J_p + W_p - U_p$  such that

$$J_p = (1 - \beta)S_p, \quad (7)$$

$$W_p - U_p = \beta S_p. \quad (8)$$

Free entry implies that the value of posting a vacancy is zero:  $V_p = 0$  for all  $p$  and, therefore,

$$\begin{aligned} c_p &= \delta q(\theta_p)E_p J_{p'} \\ &= \delta q(\theta_p)(1 - \beta)E_p S_{p'}. \end{aligned} \quad (9)$$

The Bellman equation for the surplus is:

$$\begin{aligned} S_p &= p - z + \delta(1 - s)E_p(W_{p'} + J_{p'}) + \delta E_p(sU_{p'} - f(\theta_p)W_{p'} - (1 - f(\theta_p))U_{p'}) \\ &= p - z + \delta(1 - s)E_p(W_{p'} + J_{p'} - U_{p'}) - \delta f(\theta_p)E_p(W_{p'} - U_{p'}) \\ &= p - z - \delta f(\theta_p)\beta E_p S_{p'} + \delta(1 - s)E_p(W_{p'} + J_{p'} - U_{p'}) \\ &= p - (z + \delta f(\theta_p)\beta E_p S_{p'}) + \delta(1 - s)E_p S_{p'}. \end{aligned} \quad (10)$$

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<sup>3</sup>As in Shimer (2005), we implicitly assume that the value functions depend only on  $p$  and not on  $u$ . Existence of such an equilibrium is straightforward and follows from the existence of a solution to equation 22 derived below. Uniqueness of such an equilibrium for the Pissarides (1985, 2000) model with aggregate uncertainty was proved in Mortensen and Nagypal (2005).

An existing match generates  $p$  units of output every period. It is destroyed next period with probability  $s$ . In this case, the value of the firm drops to zero, the value of a vacancy. The worker, on the other hand, becomes unemployed and gets utility  $z$  every period until he becomes employed again with probability  $f(\theta_p)$  per period. An employed worker keeps a share  $\beta$  of the match surplus. With probability  $1 - s$ , the match exists next period and generates surplus depending on the realization of  $p'$ .

We now derive the expressions for equilibrium wages and profits. Using equation 7, it follows from the free-entry condition 9 and the flow equation 3 for  $J$  that:

$$(1 - \beta)S_p = p - w_p + (1 - s)c_p/q(\theta_p). \quad (11)$$

Free entry and (10) imply that

$$S_p = p - z + (1 - s - f(\theta_p)\beta)\frac{c_p}{q(\theta_p)(1 - \beta)}. \quad (12)$$

Thus, we have that

$$(1 - \beta)S_p = (1 - \beta)(p - z) + c_p\frac{1 - s - f(\theta_p)\beta}{q(\theta_p)}. \quad (13)$$

Rearranging 11 and substituting using 13, we find that wages are given by

$$\begin{aligned} w_p &= p - (1 - \beta)S_p + (1 - s)c_p/q(\theta_p) \\ &= \beta p + (1 - \beta)z + c_p\beta\theta_p, \end{aligned} \quad (14)$$

and accounting profits are given by

$$\Pi_p = p - w_p = (1 - \beta)(p - z) - c_p\beta\theta_p. \quad (15)$$

## 2.4 Adding Capital to the Model

Our calibration strategy relies on measuring costs of posting vacancies in the data. These costs include the labor costs of time spent on hiring and costs of non-operating capital. Thus, to account for the capital costs of vacancy creation, we now follow Pissarides (2000) and add capital to the model.



We will be able to measure the labor costs of vacancy creation in the data directly. It is more difficult to measure the amount of capital residing in vacant jobs looking for workers. In the deterministic version of the model, vacancies would arise only because firms need to replace exogenously separated workers. Thus, we assume that posting firms and operating firms rent the same amount of capital. We will report the sensitivity of our results to this assumption in Section 3.4.

Let  $K$  denote the aggregate capital stock. The number of active firms equals  $v + 1 - u$ ,  $1 - u$  of them are operating and  $v$  are looking for a worker. Thus, the amount of operating capital equals  $K \frac{1-u}{v+1-u}$  and the amount of idle capital equals  $K \frac{v}{v+1-u}$ .

The aggregate constant returns to scale production function is

$$F(K \frac{1-u}{v+1-u}, A(1-u)), \quad (16)$$

where  $A$  is labor-augmenting productivity. We define  $k := \frac{K}{A(v+1-u)}$ , the capital stock per efficiency unit of labor and  $f(k) := F(k, 1)$ , the output per efficiency unit of labor. Denote by  $k^*$  the constant value of  $k$  that solves  $f'(k) = \frac{1}{\delta} - 1 + d$ , the equilibrium condition for the firm's capital stock, where  $d$  is the depreciation rate.

We can now redefine labor productivity from the preceding sections

$$p := A(f(k^*) - (\frac{1}{\delta} - 1 + d)k^*). \quad (17)$$

Because firms can buy and sell capital in a competitive market, the wage bargain is not affected and the model can be solved as before. The only difference is that  $A$ , the exogenous productivity process, is multiplied with the constant  $(f(k^*) - (\frac{1}{\delta} - 1 + d)k^*)$ . Thus,  $p$  is still an exogenous (productivity) process. The firm's flow capital cost of posting a vacancy is

$$A(\frac{1}{\delta} - 1 + d)k^*. \quad (18)$$

## 3 Cyclical Behavior of Unemployment and Vacancies

### 3.1 Calibration

In this section we calibrate the model to match U.S. labor market facts, including the cyclicity of wages and the costs of posting vacancies. The following parameters have to be



determined: average productivity  $\bar{p}$ , the value of non-market activity  $z$ , the discount factor  $\delta$ , the separation rate  $s$ , the bargaining power  $\beta$ , the vacancy cost  $c_p$ , and the matching function parameter - introduced below -  $l$ .

**Basics.** We choose the model period to be one week (one-twelfth of a quarter, to be precise), which is lower than the frequency of the employment data we use, but necessary to deal with time aggregation. The data used to compute some of the targets have monthly, quarterly or annual frequency, and we aggregate the model appropriately when matching those targets. We set  $\delta = 0.99^{1/12}$ . Shimer (2005) estimates the average monthly job finding rate from 1951 to 2003 to be 0.45 and the separation rate (not adjusted for time aggregation) to be 0.026. At weekly frequency these estimates imply a job finding rate  $f = 0.139$ , a job separation rate  $s = 0.0081$ , and a steady state unemployment rate  $u = s/(s + f) = 0.055$ .<sup>4</sup>

**Productivity.** Labor productivity,  $p$ , is measured in the data as seasonally adjusted quarterly real average output per person in the non-farm business sector constructed by the BLS from the National Income and Product Accounts and the Current Employment Statistics. This is the same measure as the one used in Shimer (2005). The stochastic process for labor productivity is chosen as follows. We approximate through a 35-state Markov chain the continuous-valued AR(1) process

$$\log p_{t+1} = \rho \cdot \log p_t + \epsilon_{t+1}, \quad (19)$$

where  $\rho \in (0, 1)$  and  $\epsilon \sim N(0, \sigma_\epsilon^2)$ . To calibrate  $\rho$  and  $\sigma_\epsilon^2$ , we consider quarterly averages of weekly productivity and HP-filter (Prescott (1986)) this process with a smoothing parameter of 1600, commonly used with quarterly data. In the data we find (similarly to Hornstein, Krusell, and Violante (2005b)) an autocorrelation of 0.765 and an unconditional standard deviation of 0.013 for the HP-filtered productivity process. At weekly frequency

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<sup>4</sup>The probability of not finding a job within a month is 0.55. The probability of not finding a job within a week then equals  $0.55^{1/4} = 0.861$  and the probability of finding a job equals  $1 - 0.861 = 0.139$ . The probability of observing someone not having a job who had a job one month ago equals (counting paths in a probability tree):  $s\{(1-f)(fs + (1-f)^2) + f(s(1-f) + (1-s)s)\} + (1-s)\{s(fs + (1-f)^2) + (1-s)(s(1-f) + (1-s)s)\} = 0.026$ . Solving for  $s$ , we obtain  $s = 0.0081$ .

this requires setting  $\rho = 0.9895$  and  $\sigma_\epsilon = 0.0034$  in the model. The mean of  $p$  is normalized to one.<sup>5</sup>

**Labor Market Tightness.** To measure the fraction  $\frac{v}{v+1-u}$  of the aggregate capital stock that is held by firms that posted vacancies but do not yet produce, we need to know the average value of  $\theta$ . Shimer (2005) estimated the average monthly job finding rate,  $f$ , to be 0.45. Den Haan, Ramey, and Watson (2000) find a monthly job filling rate,  $q$ , of 0.71. Since  $\theta = f/q$ , these numbers imply a value for  $\theta$  of  $0.45/0.71 = 0.634$ , which we choose as our calibration target. This number accords well with the direct estimate of 0.539 obtained by Hall (2005a) from the Job Openings and Labor Turnover Survey (JOLTS). As expected, this estimate is slightly lower than 0.634. JOLTS started in December 2000 and covers only a recession and a fraction of the expansion that had slower employment growth than usual. Moreover, some vacancies are not captured by JOLTS: we see firms hiring workers within a month without ever reporting having a vacancy to JOLTS.

**Matching Function.** We need a matching function that ensures that the probability of finding a job and of filling a vacancy lies between 0 and 1 (since the precise value of  $\theta$  is now meaningful, we cannot conveniently normalize it as was done in Shimer (2005)). We follow Den Haan, Ramey, and Watson (2000) (HRW) and choose

$$m(u, v) = \frac{u \cdot v}{(u^l + v^l)^{1/l}}.$$

**The Cyclicalities of Wages.** We estimate the cyclicalities of wages from BLS data (1951:1-2004:4). We find that a 1-percentage-point increase in our measure of labor productivity is associated with a 0.449-percentage-point increase in real wages. Wages are measured as labor share times labor productivity. Both time series are in logs and HP-detrended with a smoothing parameter of 1600.

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<sup>5</sup>We have defined  $p$  as the marginal product of labor. In the data we observe the average product of labor. This is inconsequential. Consider a Cobb-Douglas production function  $F(K, L) = \xi L^\gamma K^{1-\gamma}$ , and let  $p = \frac{\gamma F(K, L)}{L}$  be the marginal product of labor, and  $\hat{p} = \frac{F(K, L)}{L}$  be output per worker that we observe in the data. Note that  $\log p = \log \gamma + \log \hat{p}$ . Thus,  $\text{var}(\log p) = \text{var}(\log \hat{p})$ . Average  $p$  is then just a normalization.

To obtain the corresponding estimate in the model, we first aggregate the weekly model-generated data to replicate the quarterly frequency of the BLS data. We then log and HP-filter the time series and estimate a regression identical to the one estimated on the BLS data. The resulting regression coefficient, 0.449, is one of our calibration targets. We will discuss other estimates available in the literature and report the sensitivity of our results to them in Section 3.4.<sup>6</sup>

**The Capital Costs of Posting Vacancies.** In Section 2.4 we derived that the flow capital cost of posting vacancies equals  $(\frac{1}{\delta} - 1 + d)kA = \frac{F_K K}{v+1-u}$ , where  $F_K$  denotes the derivative of  $F$  with respect to its first argument. Decompose

$$\frac{F_K K}{v+1-u} = \frac{F_K K}{F} \frac{1-u}{1-u+v} \frac{F}{1-u}. \quad (20)$$

We now compute the steady state values for all three factors. Typical estimates from the national accounts imply a capital income share  $\frac{F_K K}{F} = 1/3$ .

Since  $\theta = 0.634$  and  $u = 0.055$ , the number of vacancies  $v = \theta u = 0.03487$ . Thus, the second factor  $\frac{1-u}{1-u+v} = 0.9644$ .

In a search model income and production shares of labor and capital do not coincide. This is because labor is paid below productivity to compensate firms for the costs of vacancy creation. However, since labor productivity is normalized to one ( $F_L A = 1$ ), it follows that

$$\begin{aligned} \frac{1-u}{F} &= \frac{F_L A(1-u)}{F} = 1 - \frac{F_K K \frac{1-u}{v+1-u}}{F} \\ &= 1 - \frac{1}{3} \frac{1-u}{1-u+v} = 1 - 0.321 = 0.679. \end{aligned}$$

Thus, the steady state capital flow cost of posting a vacancy  $c^K$  equals 0.474, or 47.4% of the average weekly labor productivity.<sup>7</sup>

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<sup>6</sup>A standard assumption of the MP model is that wages are renegotiated whenever the aggregate state of the economy changes. An alternative wage determination assumption might be that firms insure workers against aggregate income risk. It is unclear why such insurance should be provided by firms and not by financial markets, since aggregate risk is observable. We will present specific evidence justifying the standard assumption below. For now we just note that the scope of possible insurance provision is undermined by the extent of worker mobility, e.g., the average job duration of only about 2.5 years.

<sup>7</sup>One can convert these costs into units of a consumption good by dividing them by the labor share in production.

**The Labor Costs of Posting Vacancies.** The second part of the cost of filling a vacancy is the opportunity cost of labor effort devoted to hiring activities. Barron, Berger, and Black (1997) provide evidence for the time and costs involved in recruiting workers. Using the 1982 Employment Opportunity Pilot Project survey of 5700 employers, they find that on average employers spend 10.41 hours per offer and make 1.08 offers per hired worker. This implies a total of 11.24 hours spent on each hire. The corresponding numbers from the 1992 Small Business Administration survey of 3600 employers are 14.03, 1.14, and 15.99. These numbers mean that the average costs of time spent hiring one worker are between 2.2% to 3.2% of quarterly hours. Adjusting, as in Silva and Toledo (2006), for the possibility that hiring is done by supervisors who receive higher wages than a new hire, the average labor cost of hiring one worker is 3% to 4.5% of quarterly wages of a new hire. We choose the highest value of 4.5% in the benchmark calibration because this generates the lowest volatility, and we show that the results are robust to this choice in Section 3.4.

Let  $W$  be aggregate weekly wages. Wages are  $2/3$  of national income, that is,  $W = 2/3F$ . Quarterly wages then equal  $8F$ . Expected labor cost of hiring equals  $0.045 \cdot 8F$  in the data and  $c^W/q$  in the model. The probability of filling a vacancy  $q$  equals  $f/\theta = 0.219$ , and we have just found that  $F$  equals  $(1 - u)/0.679 = 1.39$ . Thus, the flow labor cost of posting a vacancy  $c^W$  equals 0.110, or 11% of the average weekly labor productivity.

**The Cyclicity of Vacancy Posting Costs.** The previous two paragraphs computed the average capital and labor costs of hiring. But both the costs of capital and wages are not constant over the business cycle.

First, capital per worker changes over the business cycle. As derived in section 2.4, firms use  $Ak^*$  units of capital in state  $A$ , where  $k^*$  solves  $f'(k) = \frac{1}{\delta} - 1 + d$ . Let  $\bar{A}$  and  $\bar{p}$  denote the mean levels of  $A$  and  $p$ , respectively. The steady state capital cost  $c^K$  then equals  $(\frac{1}{\delta} - 1 + d)k\bar{A}$  and the capital cost in state  $A$ ,  $\tilde{c}^K$ , is equal to  $c^K A/\bar{A}$ . Thus, capital cost  $\tilde{c}^K$  equals  $c^K A/\bar{A} = c^K p/\bar{p} = c^K p$  in state  $p = A(f(k^*) - (\frac{1}{\delta} - 1 + d)k^*)$  since we have normalized  $\bar{p} = 1$ .

Second, labor costs of hiring change over the business cycle according to  $c^W p^\xi$ . To determine  $\xi$  we assume that wages of those engaged in hiring are fluctuating as much

over the business cycle as do wages of other workers. As discussed above, the regression coefficient of HP-filtered log wages on HP-filtered log productivity in the data is 0.449. Since the HP-filter is a linear operator, we get the same value for  $\xi$  regardless of whether we detrend or not:  $\xi = \epsilon_{w,p} = 0.449$ .<sup>8</sup> Thus, the costs of posting a vacancy in state  $A$ , or equivalently  $p$ , equal

$$\begin{aligned} c_p &= c^K p + c^W p^{0.449} \\ &= 0.474p + 0.110p^{0.449}. \end{aligned} \tag{21}$$

**Bargaining Weights and Value of Non-market Activity.** Three parameters remain to be determined: the value of non-market activity,  $z$ , worker's bargaining weight,  $\beta$ , and the matching function parameter,  $l$ . We choose the values for these parameters to match the data on the average value for labor market tightness  $\theta = 0.634$ , the average value for the job finding rate  $f = 0.139$  and the elasticity of wages with respect to productivity  $\epsilon_{w,p} = 0.449$ . Thus, there are three targets, all described in the previous paragraphs, to pin down three parameters.<sup>9</sup>

**Computation.** First, we use the free entry condition (9) and flow equation for the surplus

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<sup>8</sup>Linearity means that  $HP(\log p^\xi) = \xi HP(\log p)$ . Thus HP-filtering an isoelastic time series does not affect the regression coefficient. Both regressions,  $HP(\log p^\xi)$  on  $HP(\log p)$  and  $\log p^\xi$  on  $\log p$ , give the same coefficient  $\xi$ .

<sup>9</sup>Note that we could have taken a different route to pin down the value of  $\beta$ . There are several papers (e.g., Christofides and Oswald (1992), Blanchflower, Oswald, and Sanfey (1996) and Hildreth and Oswald (1997)) which use cross-sectional data to test for rent-sharing in the U.S. labor market. They find that, controlling for outside labor market conditions, a one percentage point increase in profitability leads to an increase in wages of less than 0.05%. It is remarkable that their value is close to our finding of  $\beta = 0.052$ . (Since they control for our outside labor market conditions, their rent-sharing parameter corresponds to  $\beta$  in our model and not to the elasticity of wages). Note that the identification in those papers does not rely on the cyclical volatility of wages. We prefer to stay away from transplanting parameter values from one model to another, but if we were to do so, combining these estimates with our estimate of the vacancy costs would yield results identical to the ones in our paper.

(10) to derive the following difference equation in  $\theta$ :

$$\begin{aligned}
\frac{c_p}{\delta q(\theta_p)} &= (1 - \beta)E_p S_{p'} \\
&= (1 - \beta)E_p \{p' - (z + \delta f(\theta_{p'})\beta E_{p'} S_{p''}) + \delta(1 - s)E_{p'} S_{p''}\} \\
&= E_p \left\{ (1 - \beta)(p' - z) - \frac{f(\theta_{p'})\beta c_{p'}}{q(\theta_{p'})} + \frac{(1 - s)c_{p'}}{q(\theta_{p'})} \right\} \\
&= E_p \left\{ (1 - \beta)(p' - z) - c_{p'}\beta\theta_{p'} + \frac{(1 - s)c_{p'}}{q(\theta_{p'})} \right\}. \tag{22}
\end{aligned}$$

We solve this difference equation to find  $\theta$  as a function of productivity  $p$ . Next, we simulate the model to generate artificial time series for productivity, unemployment, vacancies, wages, and accounting profits. To do so, we start with an initial value for unemployment and productivity and draw a new productivity shock according to the Markov chain derived above. We then know  $\theta$  and, thus, the job finding rate and the new unemployment rate. Iterating this procedure generates the time series of interest.

The performance of the model in matching calibration targets is described in Table 1. We are able to match the targets exactly. Calibrated parameter values can be found in Table 2.

## 3.2 Main Result

The statistics of interest, computed from U.S. data, are presented in Table 3. Hornstein, Krusell, and Violante (2005b) report virtually identical numbers. Our goal in this paper is to evaluate whether a reasonably calibrated MP model can replicate these statistics.

We use the calibrated parameter values to simulate the model to create artificial time series and compute their relevant moments. Table 4 describes the results. A comparison with the corresponding statistics in the data reveals that the model matches the key business cycle facts quite well. In particular, the volatility of labor market tightness, unemployment, and vacancies is higher, but close to that in the data.<sup>10</sup>

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<sup>10</sup>Note that Table 4 reveals two shortcomings of the MP model. The correlation of labor market tightness and productivity is too high compared to the data and vacancies are more persistent in the data. This problem is well known and it arises because labor market tightness is a function of productivity only. Fujita and Ramey (2006) show that these problems are easily fixed by introducing adjustment costs in vacancy



### 3.3 Analysis

This section contains a theoretical analysis of our calibration strategy.

#### 3.3.1 Productivity Elasticity of $\theta$ and the Value of $z$

First, we consider the model without aggregate uncertainty ( $p = p'$ ) and derive the elasticity of labor market tightness with respect to aggregate productivity,  $\epsilon_{\theta,p}$ . We show that this elasticity is increasing in the value of the non-market activity of the workers,  $z$ .

In the case of no aggregate uncertainty (and constant  $c$ ) we can solve for the surplus:

$$S = \frac{p - z}{1 - \delta(1 - s) + \delta f(\theta)\beta}. \quad (23)$$

Plugging this into the free entry condition yields:

$$\frac{p - z}{1 - \delta(1 - s) + \delta f(\theta)\beta} = \frac{c}{\delta q(\theta)(1 - \beta)}, \quad (24)$$

and, equivalently,

$$\frac{1 - \delta(1 - s)}{\delta q(\theta)} + \beta\theta = \frac{p - z}{c}(1 - \beta). \quad (25)$$

Implicit differentiation delivers:

$$\frac{\partial \theta}{\partial p} = \frac{(1 - \beta)/c}{\beta - \frac{\frac{\partial q(\theta)}{\partial \theta}}{q(\theta)^2} \frac{1 - \delta(1 - s)}{\delta}} \quad (26)$$

$$= \frac{1}{p - z} \frac{\beta\theta + \frac{1 - \delta(1 - s)}{\delta q(\theta)}}{\beta - \frac{\frac{\partial q(\theta)}{\partial \theta}}{q(\theta)^2} \frac{1 - r(1 - s)}{r}} \quad (27)$$

$$= \frac{\theta}{p - z} \frac{\beta\theta q(\theta) + \frac{1 - \delta(1 - s)}{\delta}}{\beta\theta q(\theta) - \frac{\frac{\partial q(\theta)}{\partial \theta}}{q(\theta)} \frac{1 - \delta(1 - s)}{\delta}} \quad (28)$$

$$= \frac{\theta}{p - z} \frac{\beta f(\theta) + (1 - \delta(1 - s))/\delta}{\underbrace{\beta f(\theta) + (1 - \eta)(1 - \delta(1 - s))/\delta}_{k1:=}} \quad (29)$$

where  $\eta$  is the elasticity of  $f(\theta)$  with respect to  $\theta$ .

posting. Such adjustment costs generate more persistent responses of vacancies and unemployment but dampen the volatility of market tightness in the model. Our finding that the volatilities we obtain, in a model without adjustment costs, are higher than those in the data is then probably a success.

Thus,

$$\epsilon_{\theta,p} = \frac{\partial \theta}{\partial p} \frac{p}{\theta} = \frac{p}{p-z} k_1. \quad (30)$$

Given the estimated parameter values,  $k_1 = 1.43$ . Thus,  $\epsilon_{\theta,p}$  is sufficiently large only if  $p - z$  is small.<sup>11</sup> In a competitive model without search frictions  $p - z$  is zero (see footnote 2). In a search model this cannot be the case because zero operating profits would imply the absence of vacancy posting by the firms. Since we do not want to choose  $z$  arbitrarily, we use the data on vacancy posting costs and wage cyclicality to pin down  $p - z$ .

### 3.3.2 The Role of $\beta$ and $z$

In this subsection we relate the productivity elasticity of labor market tightness,  $\epsilon_{\theta,p}$ , to the level of accounting profits and the cyclical properties of wages. We show that the fact that accounting profits are small in the calibrated model (2.255% of labor productivity, or 1.55% of total revenue on average),<sup>12</sup> and wages are moderately procyclical in the data implies that the value of non-market activity,  $z$ , has to be close to the productivity level,  $p$ , and workers' bargaining weight,  $\beta$ , has to be relatively small.

Without aggregate uncertainty it holds that

$$J = p - w + \delta(1 - s)J \quad (31)$$

and, thus,

$$\begin{aligned} w &= p - (1 - \delta(1 - s))J = p - (1 - \beta)(1 - \delta(1 - s))S \\ &= p - (1 - \beta)(1 - \delta(1 - s)) \frac{p - z}{1 - \delta(1 - s) + \delta f(\theta)\beta}. \end{aligned} \quad (32)$$

Accounting profits are equal to  $p - w$ :

$$\Pi = \frac{(1 - \beta)(1 - \delta(1 - s))}{1 - \delta(1 - s) + \delta f(\theta)\beta} (p - z). \quad (33)$$

<sup>11</sup>Shimer (2005) and Costain and Reiter (2005) have also noted that  $\epsilon_{\theta,p}$  is decreasing in  $p - z$  and can be made arbitrarily large.

<sup>12</sup>Note that the relatively small accounting profits and high job finding rate suggest that the labor market search frictions are small, echoing the insight in Hornstein, Krusell, and Violante (2005a).

Finally, consider the derivative of wages with respect to productivity:

$$\frac{\partial w}{\partial p} = 1 - \frac{(1 - \beta)(1 - \delta(1 - s))}{1 - \delta(1 - s) + \delta f(\theta)\beta} \quad (34)$$

$$+ \delta\beta(1 - \beta)(1 - \delta(1 - s)) \frac{p - z}{(1 - \delta(1 - s) + \delta f(\theta)\beta)^2} \frac{\partial f(\theta)}{\partial p}. \quad (35)$$

Since  $\frac{\partial f(\theta)}{\partial p}$  is positive,  $\frac{\partial w}{\partial p}$  is small if  $\frac{(1 - \beta)(1 - \delta(1 - s))}{1 - \delta(1 - s) + \delta f(\theta)\beta}$  is large, i.e., when  $\beta$  is small. Accounting profits, on the other hand, are small only if  $(p - z) \frac{(1 - \beta)(1 - \delta(1 - s))}{1 - \delta(1 - s) + \delta f(\theta)\beta}$  is small. Thus,  $p - z$  also has to be small. The explanation is easy. Small profits mean that  $p - w$  is small, and moderately procyclical wages mean that  $w - z$  is small.

To illustrate the quantitative importance of our calibration strategy, we parameterize the model by following the common practice in the literature (e.g., Shimer (2005)) of picking  $z = 0.4$  and setting  $\beta$  equal to the steady state unemployment elasticity of the matching function, which equals  $1/(1 + \theta^{-l})$ . Whereas with a Cobb-Douglas matching function, the Hosios (1990) condition – bargaining power equals the unemployment elasticity of the matching function – can be satisfied everywhere, this is possible for HRW’s matching function only on average (because the elasticity is not constant). We determine the matching parameter  $l$  and the cost of posting vacancies  $c$  (a free and non-cyclical parameter in Shimer (2005)) to match the average job finding rate 0.139 and the average level of  $\theta = 0.634$ . We find  $l = 0.339$  and  $c = 0.995$ , which implies that  $\beta = 0.455$ . The results are summarized in Panel 1 of Table 5. As expected, the standard parameterization of the model implies that the volatility of labor market tightness, and therefore the volatility of vacancies and unemployment, is ten times smaller in the model than in the data.

We next address the claim in Shimer (2005) that Nash bargaining implies wages that are too volatile and this is why the model delivers small fluctuations of market tightness. This appears to imply that if wages were less volatile, the model would have delivered large fluctuations of market tightness. We find that this statement is not quite correct. We establish that only targeting the wage elasticity does not imply that the model will generate a high volatility of market tightness. Consider the following experiment, which chooses  $\beta$  to match the right elasticity of wages in Shimer’s parametrization (i.e., fixing  $z = 0.4$ ). We find, see Panel 6 of Table 5, that the model still generates a volatility of market tightness

that is an order of magnitude smaller than in the data. Thus, the right volatility of wages alone does not imply the right volatility of market tightness.

Since the volatility of unemployment and vacancies is quite low given the standard parameterization, we can also use a Cobb-Douglas matching function  $m(u, v) = \chi u^\alpha v^{1-\alpha}$  (with small volatilities, job finding and filling probabilities remain below one). We pick the same elasticity of the matching function as for HRW, so that  $\alpha = 0.455$ , and set  $\beta$  equal to it. We choose the matching technology parameter  $\chi$  and the cost of posting vacancies  $c$  to match the average job finding rate  $f = 0.139$  and the average level of  $\theta = 0.634$ . We find  $\chi = 0.178$  and  $c = 0.995$ . Panel 2 of Table 5 contains the results. The volatilities of labor market tightness, unemployment, and vacancies are unaffected by the choice of the matching function.

### 3.3.3 The Role of Matching Function Elasticity

The values of  $\alpha$ , the unemployment-elasticity of the matching function, used in the literature vary considerably. Shimer (2005) considers  $\alpha = 0.72$ , Andolfatto (1996)  $\alpha = 0.6$ , Farmer (2004)  $\alpha = 0.5$ , Merz (1995)  $\alpha = 0.4$  and Hall (2005a)  $\alpha = 0.235$ . Changing  $\alpha$  may have an effect on the volatility of vacancies, but it definitely has a substantial effect on the volatility of unemployment. Hall (2005a) finds the volatility of unemployment to be high, much higher than that of vacancies. In Farmer (2004) vacancies are slightly more volatile than unemployment, and in Shimer (2005) vacancies are much more volatile than unemployment. The following simple (deterministic) steady state comparative statics arguments show why. In a steady state the flows into and out of employment are equal:

$$s(1 - u) = m(u, v). \quad (36)$$

Implicit differentiation of  $u$  with respect to  $v$  gives:

$$\frac{\partial u}{\partial v} = -\frac{(1 - \alpha)m/v}{s + \alpha m/u}. \quad (37)$$

Thus, for the elasticity  $\epsilon_{u,v}$  it holds that

$$\begin{aligned} \left| \frac{1}{\epsilon_{u,v}} \right| &= \frac{s + \alpha m/u}{(1 - \alpha)m/v} \frac{1}{\theta} \\ &= \frac{\alpha}{(1 - \alpha)} + \frac{s}{(1 - \alpha)f}. \end{aligned} \quad (38)$$

As  $s$  is small relative to  $f$ , the first term  $\frac{\alpha}{(1-\alpha)}$  matters a lot. For  $\alpha = 0.72$  it equals 2.57, for  $\alpha = 0.5$  it equals 1, and for  $\alpha = 0.235$  it equals 0.31. Quite a difference.<sup>13</sup>

This difference affects the quantitative results. To illustrate this, we now parameterize the model with a Cobb-Douglas matching function by setting  $\alpha = 0.72$  (as compared to  $\alpha = 0.455$  in Section 3.3.2),  $z = 0.4$  and  $\beta = \alpha$  as in Shimer (2005). Given these choices we calibrate the matching technology parameter  $\chi$  and the cost of posting vacancies  $c$  to match the average job finding rate  $f = 0.139$  and the average level of  $\theta = 0.634$ . We find  $\chi = 0.158$  and  $c = 0.338$ . The results of this experiment are summarized in Panel 3 of Table 5. Comparing them with the results in Panel 2 of the same table, we see that, due to the increase in  $\alpha$ , the volatility of unemployment decreases substantially, whereas the volatility of vacancies actually increases. For a higher  $\alpha$ , a percentage change in vacancies results in a much smaller change in matches and, thus, unemployment. The table also shows that vacancies become more persistent. Essentially, with high  $\alpha$ , an increase in vacancies this period does not affect next period's unemployment rate very strongly. Therefore, next period's incentives to post vacancies are not very different from this period's incentives.

Viewed from a different perspective,  $var(\log v/u) = var(\log v) + var(\log u) - 2covar(\log v, \log u)$ . The discussion in the preceding paragraph implies that  $var(\log u)$  is decreasing in  $\alpha$ . Thus, for the volatility of vacancies not to increase when  $\alpha$  increases, labor market tightness has to become much less volatile. In Section 3.3.1 we have shown, for a deterministic version of the model, that the elasticity of  $\theta$  with respect to productivity equals  $\frac{1}{p-z} \frac{\beta f + (1-\delta)(1-s))/\delta}{\beta f + \alpha(1-\delta)(1-s))/\delta}$ . When  $\beta$  is relatively small, this elasticity declines substantially when  $\alpha$  is increased. Consequently, the volatilities of market tightness and unemployment are decreasing in  $\alpha$ , but the direction of the response of vacancies is ambiguous.

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<sup>13</sup>Note that despite the fact that the approximation was taken around the deterministic steady state, it is fairly accurate in the stochastic version of the model. With  $\alpha$  equal to 0.455 in our benchmark calibration, Equation 38 implies that the variance of  $\log v/u$  is close to being evenly split between (log) vacancies and (log) unemployment. This is indeed what we observe in Table 4 and Panels 1 and 2 of Table 5.

### 3.4 Robustness Checks

The standard parameterization of the MP model generates the standard deviation ( $sd$ ) of labor market tightness  $sd(\theta) = 0.013$ .<sup>14</sup> Our benchmark calibration matches the volatility of market tightness in the data and generates the  $sd(\theta) = 0.159$ . In this section we investigate the robustness of our findings to alternative measures of the costs of vacancy posting and to alternative estimates of the productivity elasticity of wages. Note that we report the results in terms of the  $sd(\theta)$  rather than the  $sd(\log \theta)$  because in some of the experiments market tightness is zero at the lowest level of aggregate productivity.  $Sd$  is meaningful because we keep the average value of  $\theta = 0.634$  constant in all the experiments.

**Labor costs of hiring.** In the benchmark calibration we have used the highest available estimate of the total labor cost of hiring a worker – 4.5% of quarterly wages of a new hire. The lowest available estimate puts this number at 3%. We now recalibrate the model with the smaller estimate, which implies a flow labor cost of posting a vacancy  $c^W = 0.073$ . The results confirm that the choice of the estimate of the labor costs of hiring has only a small impact on the results. Reducing the labor costs of vacancies increases the volatility implied by the model to  $sd(\theta) = 0.170$ .

**Capital costs of hiring.** In the benchmark we assumed that posting firms and operating firms must rent the same amount of capital. This assumption seems natural since the one-job-one-worker abstraction of the MP model precludes any reallocation of vacant capital across workers within a firm. In addition, it may not even be in a firm's interest to engage in such, presumably costly, reallocation given the job-filling rate of over 70% per month. Nevertheless, let's suppose that vacant jobs have to rent 25% more capital than producing jobs. This reduces the volatility of market tightness:  $sd(\theta) = 0.130$ . Decreasing the capital requirement of vacant jobs by 25% increases the volatility implied by the model to  $sd(\theta) = 0.205$ . These results are expected. The smaller the (capital) costs of posting vacancies, the smaller are accounting profits in the model. When profits are small, they fluctuate more

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<sup>14</sup>As usual, we HP-filter the series before computing the  $sd$  in all the experiments with smoothing parameter 1600. Not filtering the data will not alter the conclusions of this subsection at all.



in percentage terms in response to changes in productivity over the business cycle. This implies higher volatility of labor market tightness. The results indicate that to the extent that firms can rent (a fraction of) capital after a worker is found, our assumption provides an upper bound on the capital costs of vacancy creation and, thus, a lower bound on the volatilities of unemployment and vacancies in the model.

**Wage cyclicality.** Using the aggregate statistics, we found the productivity elasticity of wages  $\epsilon_{w,p} = 0.449$ , with a standard error of 0.0416. To evaluate the robustness of our results we recalibrate the model setting the cyclicality of wages at the boundary of the 95% confidence interval around this estimate. Recalibrating the model to match  $\epsilon_{w,p} = 0.367$ , holding other calibration targets fixed, we find  $sd(\theta) = 0.186$ . Recalibrating the model to match  $\epsilon_{w,p} = 0.531$ , we find  $sd(\theta) = 0.134$ . Naturally, a higher volatility of wages implies a smaller volatility of accounting profits, and, thus, a smaller volatility of market tightness. However, the performance of the model is not very sensitive to the choice of  $\epsilon_{w,p}$  in the empirically plausible range.

One can also estimate the cyclicality of wages from individual data to avoid the cyclical selection bias attributable to the entry of low wage workers into employment in booms and exit in recessions. This cyclical composition bias is important in the regression of wages on unemployment. Since we regress wages on productivity, however, both sides of the equation are affected in the same way: if workers entering in a boom are, say, 10% less productive, their wages are also 10% lower. Thus, we expect to find little difference between the estimates on the aggregate and individual data. We use the PSID data and the measures of selection-adjusted wages computed by Solon, Barsky, and Parker (1994). We again use the BLS' seasonally adjusted real average output per person in the non-farm business sector as the measure of productivity. We regress the change in log wages between each two consecutive years on the corresponding changes in log productivity and a time trend. The estimates vary, depending on the exact sample and specification used, around the value of 0.47, which is slightly higher but close to our estimate on the aggregate data.

We would also like to compare our estimate of the wage cyclicality to other ones available in the literature. It is typical in the literature (see Abraham and Haltiwanger (1995)

for a survey) to use the (un)employment rate or GNP as a cyclical indicator. We use output per worker as our measure of business cycle conditions. Our measure has two distinct advantages. First, the MP model implies that both the (un)employment rate and GNP are endogenous regressors. They, together with wages, are driven by changes in productivity and, thus, are correlated with the error term. This biases the estimated regression coefficient. We can replicate such a regression in our calibrated model. When we regress HP-filtered log wages on HP-filtered log unemployment, we obtain an estimate of  $-0.036$ . The corresponding regression in the data yields an estimate of  $-0.038$ . Similarly, when we use HP-filtered log output as a regressor, we obtain an estimate of  $0.25$ . The analogous regression of our measure of wages in the data on the BLS-constructed measure of output in the non-farm business sector yields an estimate of  $0.25$ . These numbers are consistent with those surveyed in Abraham and Haltiwanger (1995).

Second, if we were to calibrate the model to match the coefficient from a regression of wages on (un)employment or output, we would be implicitly targeting the volatility of unemployment that we want the model to account for. The reason is clear. We would not be able to match such a coefficient without forcing the model to have as high volatility of (un)employment as in the data. The procedure we use does not have this problem.

We now revisit the argument that firms may be providing insurance to workers against aggregate risk. There is little empirical support for this argument. Gomme and Greenwood (1995) and Boldrin and Horvath (1995) suggest that one may infer the presence of such insurance provision in the data from the observation that real wages are less volatile than total hours, that the labor share is not constant, and that real wages are not strongly procyclical. We note that all these observations are consistent with a search model, such as ours, that does not include any insurance contracts.

More specific evidence comes from individual data. Contracts against risk, both aggregate and idiosyncratic, have clear implications for the sequence of wages during an employment spell (Malcomson (1999)). If the firm can commit to the contract but the worker cannot, then wages are non-decreasing. If the firm cannot commit but the worker can, then wages are non-increasing. If both can commit or both cannot commit, then wages

should be constant most of the time. The findings of McLaughlin (1994) and the subsequent literature, although not concerned with contracts, imply that all three possibilities are proved to be wrong. Real wages are not constant, and there is both a large fraction of employees with real wage cuts and with real wage increases. Thus looking at individual wage sequences refutes the relevance of contracts as insurance against (aggregate) risk.<sup>15</sup>

Finally, Solon, Barsky, and Parker (1994) find that wages are similarly procyclical among those not changing employers in a given year and those switching employers. When unemployment declines by one percentage point, stayers' wages increase by 1.24% and switchers' wages increase by 1.4%, a statistically insignificant difference. Presence of insurance contracts, on the other hand, would imply that switchers' wages are substantially more cyclical.

Thus, the weight of the evidence points to a conclusion that the standard assumption of the MP model, that wages are renegotiated when aggregate productivity changes, does not appear unreasonable.

**Filtering.** We now HP-filter the model-generated time-series with a smoothing parameter of  $10^5$  as in Shimer (2005). The results are reported in Panel 4 of Table 5. Comparing them to the identical statistics in the data, reported in Panel 5 of the same table, we find that the model matches them very well.

**Proper sampling from the model-generated data.** Until now we have followed the literature in aggregating the model-generated data. To generate quarterly data we simply took averages of weekly data. This is somewhat different from the way that the statistics in the data are constructed. We now repeat the benchmark calibration while imposing sampling procedures in the model-generated data identical to those used by the BLS and the Conference Board in the data. Vacancies are measured in the last week of a month, and unemployment and employment are measured in the second week. These snapshots are

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<sup>15</sup>Beaudry and DiNardo (1991) find, using the same data set (PSID), that wages depend on the lowest unemployment rate during a job spell. This is consistent with the assumption that firms can commit to an insurance contract and workers cannot. Since the evidence on individual wage sequences speaks clearly against such contracts, the estimates likely pick up some cyclical unobserved heterogeneity, such as procyclical promotions, a conjecture confirmed for British data by Devereux and Hart (2005a,b).

our measures of monthly data in the model. Quarterly data for vacancies, unemployment, employment, and labor market tightness are still monthly averages, both in the data and in the model. Productivity is measured in the data as quarterly output divided by the quarter's employment. We compute quarterly output as the sum of weekly output. Quarterly employment is computed as the average employment of the three monthly snapshots. We recalibrate the weekly productivity process accordingly and find an autocorrelation of 0.99 and a conditional variance of 0.0033. The results of this calibration are summarized in Table 6. Comparing these results with those in Table 4, we find that they are largely unchanged.

## 4 Discussion

In this section we discuss several implications of our results. First, our estimate of the workers' bargaining power parameter is considerably lower than the unemployment elasticity of the matching function. We show that this does not mean that the economy is far from the efficient benchmark. In particular, we prove that in the presence of taxes, welfare is maximized with the worker's bargaining weight lower than that implied by the Hosios (1990) condition. Moreover, we show quantitatively that the estimate implied by our calibration procedure is very close to the efficient one. Second, we discuss our estimate that the returns to market and non-market activities are not drastically different from each other. Finally, our calibration implies strong effects of productivity shocks on unemployment. This suggests that changes to the tax or unemployment insurance policy can have non-trivial effects as well. We explore these policy implications in Section 4.3.

### 4.1 Efficiency with Taxation

So far we have abstracted from taxes. We now show that this has two consequences. First, market activity provides much higher incremental value over non-market activity than our estimate of  $z$  appears to imply. Second, the Hosios (1990) condition ceases to imply efficiency.

The crucial step when adding taxes to the model is to derive wages. They are determined

through the generalized Nash bargaining solution, which selects, for every  $p, w$  to maximize

$$(J_p(w \cdot (1 + \tau_f)))^{1-\beta} (W_p((1 - \tau_w) \cdot w) - U_p)^\beta, \quad (39)$$

where, as before,  $J_p(\cdot)$  and  $W_p(\cdot)$  are the values to the firm and the worker at productivity level  $p$ ,  $\tau_f$  is a wage tax to be paid by the firm and  $\tau_w$  is the wage tax to be paid by the worker, respectively. Set  $\tilde{w}_p = w_p(1 - \tau_w)$  and  $\hat{w}_p = w_p(1 + \tau_f)$ .

The optimum satisfies:

$$(1 - \beta)(1 + \tau_f)(W_p(\tilde{w}_p) - U_p) = \beta J_p(\hat{w}_p)(1 - \tau_w). \quad (40)$$

This implies that

$$J_p(\hat{w}_p) = (1 - \beta) \left( \frac{1 + \tau_f}{1 - \tau_w} (W_p(\tilde{w}_p) - U_p) + J_p(\hat{w}_p) \right), \quad (41)$$

and

$$W_p(\tilde{w}_p) - U_p = \beta \frac{1 - \tau_w}{1 + \tau_f} \left( \frac{1 + \tau_f}{1 - \tau_w} (W_p(\tilde{w}_p) - U_p) + J_p(\hat{w}_p) \right). \quad (42)$$

The flow equation for  $J_p$  is:

$$J_p(\hat{w}_p) = (\tilde{p} - \hat{w}_p) + \delta(1 - s)E_p J_{p'}(\hat{w}_{p'}), \quad (43)$$

where  $\tilde{p}$  is the after sales tax revenue/productivity.

The free entry equation is unchanged:

$$c_p = \delta q(\theta_p) E_p J_{p'}(\hat{w}_{p'}). \quad (44)$$

Define the surplus

$$S_p = \left( \frac{1 + \tau_f}{1 - \tau_w} (W_p(\tilde{w}_p) - U_p) + J_p(\hat{w}_p) \right). \quad (45)$$

It holds that

$$\begin{aligned} S_p &= \frac{1 + \tau_f}{1 - \tau_w} (\tilde{w} + \delta(1 - s)E_p(W_{p'}(\tilde{w}_{p'})) + \delta s E_p(U_{p'})) \\ &\quad - \frac{1 + \tau_f}{1 - \tau_w} (z + \delta f(\theta_p)E_p W_{p'}(\tilde{w}_{p'}) + \delta(1 - f(\theta_p))E_p U_{p'}) \\ &\quad + (\tilde{p} - \hat{w}) + \delta(1 - s)E_p J_{p'}(\hat{w}_{p'}) \\ &= \tilde{p} - \frac{1 + \tau_f}{1 - \tau_w} z + \delta(1 - s)E_p S_{p'} - \delta f(\theta_p)\beta E_p S_{p'}. \end{aligned} \quad (46)$$

Thus,

$$J_p(\hat{w}) = (1 - \beta)(\tilde{p} - \frac{1 + \tau_f}{1 - \tau_w}z) + \delta(1 - s)E_p J_{p'}(\hat{w}) - c_p \beta \theta_p. \quad (47)$$

For wages we have that

$$(\tilde{p} - \hat{w}_p) = (1 - \beta)(\tilde{p} - \frac{1 + \tau_f}{1 - \tau_w}z) - c_p \beta \theta_p, \quad (48)$$

and

$$\hat{w}_p = \beta \tilde{p} + (1 - \beta) \frac{1 + \tau_f}{1 - \tau_w}z + c_p \beta \theta_p, \quad (49)$$

and

$$\tilde{w}_p = \beta \frac{1 - \tau_w}{1 + \tau_f} \tilde{p} + (1 - \beta)z + c_p \beta \frac{1 - \tau_w}{1 + \tau_f} \theta_p. \quad (50)$$

Profits equal

$$\Pi = \tilde{p} - \hat{w}_p = (1 - \beta)\tilde{p} - (1 - \beta) \frac{1 + \tau_f}{1 - \tau_w}z - c_p \beta \theta_p. \quad (51)$$

Effective average tax rates, which are consistent with the concept of aggregate tax rates at the national level, are provided by Mendoza, Razin, and Tesar (1994). In 1987 the consumption tax rate equaled 5.1% and the labor tax rate equaled 29.1%.<sup>16</sup> Their results imply  $\tau_f = 0$ ,  $\tau_w = 0.291$  and  $\tilde{p} = (1 - 0.051)p$ .

The above equations (free entry condition, solution for wages, etc.) are, given our calibration strategy, identical in the model with and without taxes. Thus, the presence of taxes does not affect the dynamics of the endogenous variables, such as market tightness and unemployment and there is no need to recalibrate and recompute the model. However, when we estimate  $z$ , we really estimate  $\frac{1 + \tau_f}{1 - \tau_w}z$ . Our estimate for  $z$  is 0.955 but the true value of  $z$  is 0.677. Instead of normalizing  $p$  to 1 we really normalize  $\tilde{p}$  to be 1. The implicit normalization on  $p$  is then  $p = 1/0.949 = 1.054$ .<sup>17</sup> Thus,  $p - z = 0.375$ .

<sup>16</sup>Lucas (1990) and Prescott (2004) use an effective tax rate of 0.4. The reason for their higher value is that in a competitive model the marginal tax rate matters, whereas here the average tax rate is relevant.

<sup>17</sup> This calculation implicitly assumes that unemployed workers do not pay consumption tax on  $z$ . This would be true if  $z$  represented only the value of leisure. Under the alternative assumption that the consumption of  $z$  is fully taxed, consumption taxes do not create a wedge between the values of market and non-market activities. Therefore, we can ignore them and have  $\tilde{p} = p$ . In this case  $p - z = 0.323$ .



A second implication of taking into account taxation is that we can compute the bargaining power that maximizes social welfare. This number will be lower than the unemployment elasticity of the matching function.

The efficient levels of  $\theta$ 's are the solution to the following optimization problem:

$$SW_p(u) = \text{Max}_\theta (zu + p(1 - u) - cu\theta + \delta E_p SW_{p'}(s + (1 - s)u - f(\theta)u)). \quad (52)$$

Using the envelope theorem ( $\delta E_p SW_{p'}' = -c/f'(\theta_p^*)$ ) and the first order condition for  $\theta$ , one can show that the optimum for productivity level  $p$ ,  $\theta_p^*$ , satisfies:<sup>18</sup>

$$\frac{c}{\delta f'(\theta_p^*)} = E_p \left\{ (p' - z) + c(\theta_{p'}^* - \frac{f(\theta_{p'}^*)}{f'(\theta_{p'}^*)} + \frac{(1 - s)}{f'(\theta_{p'}^*)}) \right\}. \quad (53)$$

Consider a deterministic version and let  $\theta^*$  be the optimal market tightness. It solves

$$\frac{c}{\delta f'(\theta^*)} = (p - z) + c(\theta^* - \frac{f(\theta^*)}{f'(\theta^*)} + \frac{(1 - s)}{f'(\theta^*)}). \quad (54)$$

For  $\delta = 0.9992$ ,  $s = 0.0081$ ,  $c = 0.584$ ,  $p = 1.054$ ,  $z = 0.677$  and  $l = 0.407$ , we find  $\theta^* = 0.670$ .

We can now solve for the bargaining power such that the efficient amount of vacancies is posted. Analogously to the derivation of equation 22 we use the flow equation 46 and the free entry condition 44 to derive the equation that determines the labor market tightness for a given bargaining power of a worker in a deterministic version of the model:

$$\frac{c}{\delta q(\theta^*)} - \frac{(1 - s)c}{q(\theta^*)} = (1 - \beta)(\tilde{p} - \frac{1 + \tau_f}{1 - \tau_w}z) - c\beta\theta^*. \quad (55)$$

The result is  $\beta = 0.152$ . If the consumption of  $z$  is taxed as well (see footnote 17), we would find  $\theta^* = 0.596$ , and efficient  $\beta = 0.056$ .

This result means that our calibration strategy implies that the model is much closer to the efficient benchmark than what is implied by the standard calibration, which, paradoxically, is targeting efficiency. We see this result as a warning against calibrating the model to an efficiency condition. The efficiency condition in the MP model is not robust

<sup>18</sup>For a Cobb-Douglas matching function this simplifies to  $\frac{c}{\delta q(\theta_p^*)} = E_p \{ (1 - \alpha)(p' - z) - c\alpha\theta_{p'}^* + \frac{(1 - s)c}{q(\theta_{p'}^*)} \}$ , where  $\alpha$  denotes the unemployment elasticity of the matching function.

to small changes in the modeling environment. In addition, the purpose of calibration and quantitative analysis is to account for the world we live in, not the efficient world we may want to live in.

## 4.2 The Value of Non-Market Activity

Our calibrated value for  $z$ , the value of non-market activity, is substantially higher than the value used in the standard parameterization of the MP model. Our estimate appears reasonable since  $z$  is a sum of the value of leisure, unemployment benefits, home production, self-employment, dis-utility of work, etc. A value of  $z \approx 0.4$ , typically used in the literature, on the other hand, seems to be unreasonably low, as non-market activity is identified with receiving unemployment benefits only.

The large and strongly procyclical flows from out-of-the-labor-force into employment suggest that the value of not working has to be close to the value of working for these individuals. Otherwise, it seems hard to rationalize these flows, given the relatively small changes of productivity over the business cycle. In addition, Anderson and Meyer (1997) and Vroman (2002) document strikingly low reciprocity rates for unemployment insurance benefits among unemployed workers in the U.S. A substantial fraction of the unemployed is ineligible. The reasons for ineligibility are revealing about the value of  $z$ . The first major reason for ineligibility is that workers quit their last job (suggesting that the value of working may well be below the value of becoming unemployed). The second main reason for ineligibility is not working enough, i.e., very sporadic work patterns (suggesting weak labor force attachment of these workers and consistent with large flows in and out of out-of-the-labor-force).<sup>19</sup>

Our finding that a typical unemployed worker does not suffer a large decline in utility has to be interpreted with caution, however. In the model,  $z$  does not depend on the length of the unemployment spell. This is a strong assumption. Long-term unemployed are definitely

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<sup>19</sup>Among eligible workers, applications rates are low as well. The reasons are telling: 37% do not bother to apply because they expect to find a job fast; 7% - “too much hassle to apply,” 6% - “too much like welfare,” 6% - “do not know about UI,” 17% - “don’t know,” etc.

worse off as they face problems replacing their durable consumption goods (a broken TV, dishwasher, microwave, etc.). Furthermore, having a month off to enjoy leisure has a high value, but the enjoyment of a year of unemployment is questionable. In our calibration we (implicitly) estimate the average  $z$  of all unemployed. Since the job finding rate equals 45% per month on average, short-term unemployed make up the bulk of observations. Thus, our estimate of  $z$  represents the value of unemployment for the average worker, who finds employment quickly. It is not informative about the value of long-term unemployment, since this is a low probability event.

Allowing  $z$  to decrease with the length of the unemployment spell makes  $z$  endogenous. When productivity declines, the average duration of unemployment increases and thus the average  $z$  of the unemployment pool declines as well. This is an interesting and, we believe, productive way to add curvature to the model on the worker side. Modeling curvature in this way is unlikely to dampen the model's ability to replicate business cycle facts. It creates some procyclicality in  $z$ , but our calibration strategy would then reduce the bargaining power to match the cyclicalities of wages.<sup>20</sup> Since the effects of productivity shocks are relatively short-lived, the average duration of unemployment and, thus, the average  $z$  are unlikely to change much over the business cycle.<sup>21</sup>

It is also well documented that the consumption of workers who become unemployed exhibits only modest changes. Gruber (1997) reports that becoming unemployed is associated with a 6.8% decline in consumption expenditure. Aguiar and Hurst (2005) make a convincing case that actual consumption declines much less than consumption expenditure because unemployed workers increase the time spent shopping, looking for bargains, and preparing food at home. Measuring actual consumption, they document that the food consumption of

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<sup>20</sup>The value of non-market activity can be cyclical for several other reasons, such as correlation between market and home technology. Unfortunately, the correlation between  $p$  and  $z$  and the bargaining power  $\beta$  are not separately identified from the observation of wages  $w_p = \beta p + (1 - \beta)z + c\beta\theta_p$ . A related problem plagues the estimation of real business cycle models with home production (e.g., McGrattan, Rogerson, and Wright (1997)). They cannot simultaneously identify the elasticity of substitution between market and non-market activities and the correlation between market and non-market productivity.

<sup>21</sup>Our discussion here is reduced form. This is intentional. The structural model of the evolution of  $z$  and its identification are too complex to be adequately developed in this paper.

unemployed workers is 5% lower than the food consumption of employed workers. Even this is likely to be an upper bound because this estimate is obtained in a cross-sectional data set. If unemployed workers are selected from a non-random subsample with lower income and consumption on average, the *drop* in their consumption may be much lower. This argument is supported by the results in Stephens (2001), who finds, for the sample of workers displaced due to plant closings, that the full fall in consumption is realized over several years *prior* to the displacement. Gruber (1997) also reports that unemployed workers have a lower real wage and consumption level prior to becoming unemployed.

Our finding of a high value for  $z$  implies that the value of being unemployed is close to the value of working. Indeed, even the parameterization in Shimer (2005) implies that  $(W - U)/W \approx .003$ , i.e., that the value of becoming unemployed is just three tenths of 1% lower than the value of working. Our finding, however, does not rule out that becoming unemployed can cause noticeable distress for some individuals. Indeed, as Jacobson, LaLonde, and Sullivan (1993) find, some displaced workers suffer a substantial decline in post-displacement earnings and often go through a long period of unemployment. We note, however, that this is not caused by the search frictions of the MP model. It is probably caused by the loss of the worker's union status or the loss in the value of the worker's occupation-specific human capital. In other words, it is caused by a low post-displacement  $p$  and not by a low  $z$ .

### 4.3 Response of the Model to Changes in Policy

Costain and Reiter (2005) suggest that  $z$  cannot be too large relative to market productivity in the MP model because in this case changes in unemployment insurance would have strong effects on unemployment. They suggest that these effects are counterfactual. Unfortunately, the effects of changes in unemployment insurance are hard to measure in the data. There exist a number of microeconomic studies, surveyed in Meyer (1995), that suggest that these effects are small. A problem with this approach to evaluating the MP model lies in the fact that, in general, a change in policy affects two decisions: those by the unemployed to search and those by firms to post vacancies. Microeconomic studies address the first decision.

In a typical microeconomic study, a small fraction of the unemployed are given a bonus if they find a job fast. Their expected duration of unemployment does not decrease by too much. This is just what the MP model predicts. When we replicate these studies in the model, we find a very small effect on duration of unemployment. Because matching is random, firms' expected profits do not change when a small fraction of the unemployed has a higher  $z$ . Thus, their vacancy posting decisions are virtually unaffected. Of course, the effect of a permanent subsidy to all unemployed in the MP model would be substantial because of the response of vacancy posting decisions.<sup>22</sup> However, this is not the effect that microeconomic studies are designed to measure. In our view, microeconomic studies explore the elasticity of search decisions to monetary incentives. We take the evidence that these effects are small as a justification for not endogenizing a worker's search effort.

We are also skeptical that the macro effects of unemployment insurance can be isolated and that endogeneity problems can be overcome in cross-country regressions. What matters for the MP model is not the level of unemployment insurance per se, but the size of the match surplus that incorporates the effects of unemployment insurance, tax structure (e.g., capital vs. labor vs. consumption taxes that have very different effects on the match surplus, in particular, depending on the extent to which the consumption of  $z$  is taxed), subsidies to firms, subsidies to workers (e.g., subsidized childcare), etc. In addition, an increase in unemployment insurance does not increase  $z$  one-for-one. For example, Gruber and Cullen (2000) find that for each dollar of a husband's unemployment insurance received, wives earn 73 cents less. Moreover, a higher replacement rate crowds out private (precautionary) savings (Gruber and Engen (2001)). Taking into account the latter two effects is important to quantitatively assess the effect of changes in unemployment insurance in a cross-country comparison since the degree of women's labor force participation and the development of financial markets differ between countries. But all these important determinants of labor supply are not held constant in cross-country regressions of the unemployment rate on

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<sup>22</sup>Interestingly, Vroman (2002) documents that, even in the aggregate, only just over a quarter of all unemployed receive unemployment insurance benefits. Thus, holding everything else constant, a change in unemployment insurance affects the wages of about 25% of the unemployed. As a result, expected profits decrease by much less than in a world where all unemployed receive unemployment insurance.

unemployment benefits (e.g., surveyed in Layard and Nickell (1999)). Thus, we think that results from such regressions cannot be used to assess the success of a model.<sup>23</sup>

Any model where shocks to productivity are strongly amplified is likely to exhibit strong effects of policies as well. The argument is simple. Any sequence of productivity shocks can be replicated through a sequence of sales taxes. In a basic real business cycle model, productivity and tax changes have identical effects both on first-order conditions and on households' budget constraint – the conditions that characterize the equilibrium. In a model with employment lotteries (e.g., Rogerson (1988)), increasing unemployment insurance acts like a wage tax, the effects of which are close to those of a productivity shock. Thus, with our calibration strategy, the MP model joins the set of widely used models that can rationalize differences in output and unemployment through differences in policy.<sup>24</sup>

Our major concern with the policy analysis in the MP model lies in its linearity. We define  $p = F_L$ , a process that changes with changes in technology, in capital and in employment. The variation of employment and capital over the business cycles creates some curvature in  $p$ , which is absent in our analysis since we take  $p$  to be an exogenous process. This is fine for our purposes in this paper, because what matters is how much  $p$  varies over the business cycles (measured in the data) and not whether technology, capital or employment cause this movement. It has to be recognized, however, that, with curvature in labor in production, one cannot treat  $p$  as an exogenous process when studying the effects of changes in policy, especially if large changes in the employment level are considered. In other words, with a decreasing marginal product of labor,  $p$ , a 1-percentage-point increase in  $p$  requires more than a 1% change in technology. This does not affect our results on the amplification of productivity, since what matters for our analysis in this paper and what we measure in the data is  $p$ . But the effect of a change in, say, tax rates is dampened through

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<sup>23</sup>More direct evidence on economy-wide changes in policy is provided by studies on the effects of the earned income tax credit (EITC). The consensus in the literature (summarized in Hotz, Mullin, and Scholz (2001)) seems to imply very strong effects of changes in this policy on employment.

<sup>24</sup>Prescott (2004), for example, finds that differences in marginal tax rates alone can explain why Americans work 50% more than do the Germans, French, and Italians. And Ljungqvist and Sargent (1998) argue that the welfare state is responsible for the persistently high European unemployment rates.

the curvature in labor.

Adding curvature in the value of non-market activity also dampens the effects of policies. Recall our discussion that  $z$  is likely to be decreasing with the length of the unemployment spell, for example, as in the top panel of Figure ???. Consider a change in policy, such as an increase in tax rates or unemployment benefits, that increases  $z$  relative to  $p$ . A stylized illustration of the effects is illustrated in the bottom panel of Figure ???. In response to such a policy firms post fewer vacancies. This leads to an increase in the average duration of unemployment accompanied by a decline in the average  $z$  of the unemployment pool. This works against the direct effect of the policy and moves the economy closer to the equilibrium prior to the change in the policy. As discussed above, this is unlikely to dampen the model's ability to replicate business cycle facts because cyclical fluctuations in productivity are relatively short-lived as compared to more permanent changes in policy. Depending on the curvature of  $z$ , however, the policy's effect may be entirely canceled out.

The size of changes in unemployment insurance is presumably dampened further if we allow for on-the-job-search as in, say, Pissarides (1994) or Burdett and Mortensen (1998). Firms post vacancies to attract both employed and unemployed workers, but  $z$  has a strong effect on the wages of the unemployed only. The overall effect of  $z$  on profits and vacancy posting depends on how much profits firms make by hiring an employed worker versus hiring an unemployed one. If firms prefer to hire the employed worker, as in Nagypal (2004), the effect of changes in  $z$  will be reduced.

Another feature of the labor market that diminishes the effect of changes in unemployment insurance is the presence of minimum wages and of the public sector with largely administered wages. For some workers, a government job is the relevant outside option; for others, the minimum wage is binding, i.e., it is higher than the outcome of negotiations  $w_p$ . Changes in unemployment insurance then affect wages only of the remaining workers, whose wages equal  $w_p$  and whose outside option is unemployment (if the minimum wage is still binding for others). Adding these institutional features to the model does not affect the volatility of  $v$  or  $\theta$ , since the model has to be recalibrated. As before, the elasticity of  $v$  and  $\theta$  depends on the size of accounting profits (implied by the vacancy posting costs)

and the average elasticity of wages, our calibration targets. Of course, calibrated parameter values will change, but the elasticity and, thus, the volatility will not.

## 5 Conclusion

We have proposed a new way to calibrate the parameters of the Mortensen-Pissarides model and found that a reasonably calibrated model is consistent with the key business cycle facts. In particular, it generates volatilities of its key variables - unemployment, vacancies, and labor market tightness - that are very close to those observed in the data. Given the recent controversy in the literature on this issue, and the popularity of the MP model in quantitative work, our finding is comforting.

We explain why the standard parametrization fails to generate right volatilities in the model. This is due to placing assumptions on the values of the key parameters - the value of non-market activity and the bargaining weight of the workers - that are inconsistent with the data. We use data on the costs of posting a vacancy and on the cyclicalities of wages to identify these parameters.

Our calibration implies that the value of non-market activity is fairly close to market productivity. This is the result one would expect in a frictionless competitive environment. It then seems reasonable that a search and matching model, which shares many features with an RBC model, exhibits a similar relationship. A typical practice in calibrating the MP model is to set the value of non-market activity to the replacement rate of unemployment insurance. We note, however, that it also includes the value of leisure and home production and, given the structure of the model, the costs of earning income. We should reiterate that we found a high value of non-market activity only for the average worker, i.e., a worker with an expected unemployment duration of about two months. Our finding does not imply a high value of non-market activity for the long-term unemployed. We simply do not observe many such individuals in the unemployment pool.

We also find a relatively low value for workers' bargaining weight. We show that such a low bargaining weight is needed to restore efficiency in the MP model, once we account for the level of taxes observed in the data. Despite the apparently low bargaining weight,



however, worker's bargaining position is not weak because outside opportunities have significant effects in a dynamic model. Thus, the low bargaining weight per se does not imply that wages are either substantially below the marginal product or that wages do not change with changes in productivity.

Our finding that the MP model matches the business cycle facts very well suggests that it might be useful to study the effects of policies. A key shortcoming of the standard model for the purpose of policy analysis is its linearity. This is easily fixed through embedding the MP model into the RBC framework. As Merz (1995) and Andolfatto (1996) have shown, this significantly improves the performance of the real business cycle model as well. An incomplete list of successes includes the findings that productivity leads total hours, unemployment and vacancies are negatively correlated (Beveridge curve), and total hours and output fluctuate substantially more than wages. But the RBC model (with MP embedded and calibrated in the standard way) exhibits the same empirical shortcoming as the MP model itself. Unemployment and vacancies are not volatile enough. Applying our calibration strategy within an RBC framework resolves this problem.

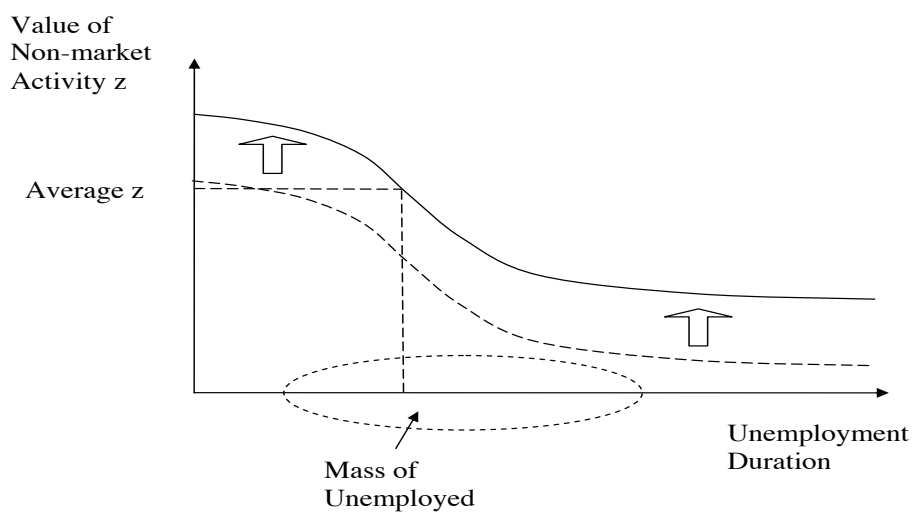
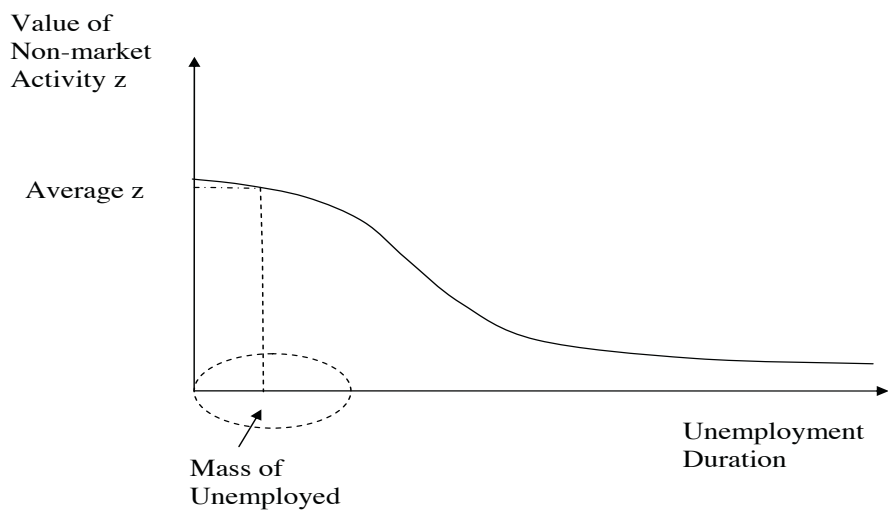


Table 1: Matching the Calibration Targets.

Target	Value	
	Data	Model
1. Elasticity of wages w.r.t. productivity, $\epsilon_{w,p}$ ,	0.449	0.449
2. Average job finding rate, $f$ ,	0.139	0.139
3. Average market tightness, $\theta$ ,	0.634	0.634

Note - The table describes the performance of the model in matching the calibration targets.

Table 2: Calibrated Parameter Values.

Parameter	Definition	Value
$z$	value of non-market activity	0.955
$\beta$	workers' bargaining power	0.052
$l$	matching parameter	0.407
$c$	cost of vacancy when $p = 1$	0.584
$\delta$	discount rate	$0.99^{1/12}$
$s$	separation rate	0.0081
$\rho$	persistence of productivity process	0.9895
$\sigma_\epsilon^2$	variance of innovations in productivity process	0.0034

Note - The table contains the calibrated parameter values in the benchmark calibration.

Table 3: Summary Statistics, quarterly U.S. data, 1951:1 to 2004:4.

		$u$	$v$	$v/u$	$p$
Standard Deviation		0.125	0.139	0.259	0.013
Quarterly Autocorrelation		0.870	0.904	0.896	0.765
Correlation Matrix	$u$	1	-0.919	-0.977	-0.302
	$v$	—	1	0.982	0.460
	$v/u$	—	—	1	0.393
	$p$	—	—	—	1

Note - Seasonally adjusted unemployment,  $u$ , is constructed by the Bureau of Labor Statistics (BLS) from the Current Population Survey (CPS). The seasonally adjusted help-wanted advertising index,  $v$ , is constructed by the Conference Board. Both  $u$  and  $v$  are quarterly averages of monthly series. Average labor productivity  $p$  is seasonally adjusted real average output per person in the non-farm business sector, constructed by the BLS from the National Income and Product Accounts and the Current Employment Statistics. All variables are reported in logs as deviations from an HP trend with smoothing parameter 1600.

Table 4: Results from the Calibrated Model.

		$u$	$v$	$v/u$	$p$
Standard Deviation		0.145	0.169	0.292	0.013
Quarterly Autocorrelation		0.830	0.575	0.751	0.765
Correlation Matrix	$u$	1	-0.724	-0.916	-0.892
	$v$	—	1	0.940	0.904
	$v/u$	—	—	1	0.967
	$p$	—	—	—	1

Note - All variables are reported in logs as deviations from an HP trend with smoothing parameter 1600. Calibrated parameter values are described in Table 2.

Table 5: Experiments.

	$u$	$v$	$v/u$
1. <i>HRW</i> , $l = 0.399$ , $\beta = 0.455$ , $z = 0.4$			
Standard Deviation	0.010	0.012	0.022
Quarterly Autocorrelation	0.827	0.610	0.762
2. <i>Cobb-Douglas</i> , $\alpha = \beta = 0.455$ , $z = 0.4$			
Standard Deviation	0.010	0.012	0.022
Quarterly Autocorrelation	0.824	0.612	0.760
3. <i>Cobb-Douglas</i> , $\alpha = \beta = 0.72$ , $z = 0.4$			
Standard Deviation	0.005	0.016	0.021
Quarterly Autocorrelation	0.827	0.715	0.763
4. <i>Our Calibration</i> , <i>HP-Filter Smoothing Parameter</i> $10^5$			
Standard Deviation	0.194	0.204	0.380
Quarterly Autocorrelation	0.901	0.713	0.851
5. <i>Data</i> , <i>HP-Filter Smoothing Parameter</i> $10^5$			
Standard Deviation	0.190	0.202	0.382
Quarterly Autocorrelation	0.936	0.940	0.941
6. <i>HRW</i> , $z = 0.4$ , $\epsilon_{w,p} = 0.449$			
Standard Deviation	0.010	0.012	0.021
Quarterly Autocorrelation	0.826	0.610	0.761

Note - Panel 1 replicates the experiment in Shimer (2005) but with the HRW-matching function and our calibrated value of  $l = 0.399$ . Panel 2 replicates the experiment in Shimer (2005) with Cobb-Douglas matching function and setting  $\alpha = 0.455$  – the theoretical matching function elasticity from Panel 1. Panel 3 replicates the experiment in Shimer (2005) with Cobb-Douglas matching function and  $\alpha = 0.72$ . Panel 4 reports the results of our calibration when all the variables in the calibrated model are HP-filtered with smoothing parameter  $10^5$ . Panel 5 reports the corresponding statistics in U.S. data. Panel 6 replicates the experiment in Shimer (2005) but with the additional target  $\epsilon_{w,p} = 0.449$ .

Table 6: Calibration Results with Proper Sampling from the Model-Generated Data.

		$u$	$v$	$v/u$	$p$
Standard Deviation		0.148	0.175	0.293	0.013
Quarterly Autocorrelation		0.831	0.554	0.765	0.765
Correlation Matrix	$u$	1	-0.830	-0.956	-0.941
	$v$	—	1	0.957	0.929
	$v/u$	—	—	1	0.977
	$p$	—	—	—	1

Note - All variables are reported in logs as deviations from an HP trend with smoothing parameter 1600.

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